Local Labor Markets in Mexico

by

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ABSTRACT

This dissertation studies how shocks to regional economies in Mexico shape employment, wages, human capital, fertility, and marriage in the short and medium run.

The first chapter studies the causal impact of fluctuations in men's and women's labor market opportunities on fertility. I evaluate how jobs in the formal sector, in manufacturing, and at export-assembly plants (maquiladoras) in Mexico shape childbirth, selection into fertility, and the timing of births using two complementary identification strategies. The first strategy exploits exogenous shocks to demand for male versus female labor using a region's industrial structure, and the second uses establishment-level data from the universe of maquiladoras to construct an instrumental variable based on large expansions and contractions in plant employment. Results show that positive shocks in the short run to men's employment have large, positive effects on fertility, whereas positive shocks to women's employment have small net impacts in the short run.

The second chapter tests the predictions of traditional models of the family. These models assume that men specialize in market work, whereas women specialize in household production. An implication of these models is that increases in women's wages, relative to men's wages, decrease the gains to marriage. Previous work has struggled with generating an accurate proxy for women's potential wages in middle-income countries, where women's

labor force participation is often low and subject to selection bias. I create a measure of women's potential wages in the market, which applies regardless if women actually work, and show that the predictions of the models hold in Mexico. I also find that women are more likely to be heads of households and to be single mothers in response to increases in their relative wages.

The third chapter evaluates the causal impact of Chinese export shocks on Mexico's local labor markets. The findings indicate that important margins of adjustment to labor demand shocks in Mexico differ from those found in other studies on wealthy countries. Municipalities with greater exposure to Chinese trade penetration do not experience bigger drops in the share of the population employed in manufacturing, nor are other measures on the extensive margin of employment affected. Instead, I find large negative impacts on wages, human capital levels, and the skilled-labor share of manufacturing. I also find consistently negative impacts across the conditional wage distribution, with workers in higher quantiles in manufacturing suffering slightly larger wage decreases than workers in lower quantiles.

CHAPTER I

Fathers and Sons: Labor Market Opportunities and Fertility

1.1 Introduction

This paper investigates how annual fluctuations in labor markets in Mexico for men and women affect decisions about fertility. In the last few decades, many developing countries have industrialized, and large numbers of women have entered the labor force for the first time. Much of the rise in women's employment is due to the growth of export-oriented manufacturing employment, especially in low-skill jobs. This study contributes to a recent, growing literature evaluating the consequences of this remarkable transition in women's labor force attachment on families' decisions about fertility and marriage. Unlike other work that focuses solely on positive shocks to women's labor market opportunities (Heath and Mubarak, 2015; Sivasankaran, 2014; Jensen, 2012), however, I study both expansions and contractions in employment for men and women and show that labor market shocks for the former also matter in fertility decisions.

I exploit multiple sources of data over different time periods to build two complementary instrumental variables identification strategies to study how fluctuations in labor market opportunities in the formal sector for both men and women affect fertility, selection into fertility, and the timing of births. The first approach generates a predicted measure of employment, building on an approach commonly used in labor and urban economics to isolate exogenous shocks to labor demand, to identify how these local labor demand shocks shift fertility. The second approach uses a restricted-access, establishment-level dataset on the universe of maquiladoras (export-assembly plants) to construct an instrument based on expansions and contractions at the factory-level.

I find that increases in labor market opportunities for men have large, positive impacts on fertility, whereas increases in women's labor market opportunities show negligible impacts on total fertility. These results are consistent with neoclassical theories of fertility dating back to Becker (1960): families choose whether to have a child in each period, and changes in employment for men and women generate income and substitution effects that alter the proportion of households choosing to have a child.

This paper also evaluates how employment dynamics impact fertility. Distributed lag models provide evidence that the immediate effect of demand shocks for women's employment is negative, indicating that substitution effects dominate income effects in the near term. The net effect in the long run, however, is close to zero, indicating that employment shocks for women are not associated with large changes in total fertility.

I then decompose the variation in employment shocks into lower-frequency and higher-frequency components. I find that high-frequency negative demand shocks to women's labor increase fertility. As high-frequency movements in employment reflect transitory, unanticipated changes in labor market opportunities, it is likely that these accelerate the timing of fertility as families take advantage of the reduced opportunity cost of having children (Heckman and Walker, 1990). On the other hand, I find that low-frequency movements in women's labor market opportunities are not associated with significant impacts on fertility. This is consistent with income and substitution effects roughly offsetting each other in the long run.

The relationship between female labor force participation and fertility remains a matter of debate (Engelhardt and Prskawetz, 2004; Kögel, 2004; Mishra, Nielsen, and Smyth, 2010).

Lim (2009) notes that the cross-sectional relationship within countries has become much less steep across time, and in some regions there is no relationship at all.[1] If women's opportunity cost of time, measured by wages or hours worked, is a central mechanism driving fertility decisions, as in neoclassical models following Becker (1960), then why is the relationship not robust across time and space?[2]

That theory does not provide a firm prediction about how changes in labor markets should affect fertility may explain some of the differing results across studies. The empirical analysis in this paper is informed by two points. First, labor demand shocks may lead to heterogeneous effects that differ across households, depending on initial labor force participation, labor market frictions (Da Rocha and Fuster, 2006), availability of childcare arrangements (Del Boca, 2002), other policies that alter the incentives to have children, or the quality of the work itself.[3]

Second, the interaction between labor markets for men and women is critical. The bulk of

[1] She shows that there is no clear pattern in Asia-Pacific countries, while both female labor force participation and fertility fell in many countries in the Middle East and North Africa in the 1990s.

[2] A neoclassical labor supply model predicts income and substitution effects resulting from increased wages. There is no theoretical reason to assume that substitution effects will dominate income effects for women, though existing research focusing on both men's and women's labor market opportunities (e.g. Schultz, 1985; Schaller, 2016) supports the view that substitution effects dominate. Typically, it is thought that at low wages, increases in wages lead to more hours worked as the substitution effect of higher wages dominates the income effect. This results in less time for leisure and raises the cost of having a child. At higher wages, increases in the wage may lead to reduced hours of work if the income effect dominates the substitution effect. Of course, as noted by Ahn and Mira (2002), in reality most individuals do not pick their optimal hours of work. Imagine, for simplicity, that work options are binary: an individual may either accept a full-time job or not. If she is already working, then an exogenous shock to her wage, say through a positive labor demand shock, induces only an income effect. That implies that an increase in a woman's wage would lead to increased fertility, assuming children are normal goods. Lindo (2010) and Black et al. (2013) provide evidence that children are indeed normal goods.

[3] Employment is not merely associated with changes in wages and the cost of time. Lim (2009) argues that "increases in labor force participation have not been matched by improvements in job quality and that the kinds of jobs women are engaged in and their working conditions have not led to their true socio-economic empowerment, have not provided adequately satisfying alternatives to childbearing or have not involved serious incompatibility between paid and unpaid work." She suggests a number of mechanisms linked to women's employment that should lower fertility, such as whether the quality of the work enhances women's status, thereby increasing their independence and bargaining power within the household. Other research (Ñopo, 2012; World Bank, 2011) documents that many women's jobs in Latin America are low-wage and highly segregated from men's jobs, implying that employment opportunities may not sufficiently lift women's status to change fertility decisions. The findings in this paper show that while labor demand shocks to men's labor increase wages, labor demand shocks to women's labor actually lead wages to decline. Part of this result, however, may be due to changes in the composition of the labor force.

the research on fertility focuses exclusively on either male wages or female wages, aggregate unemployment, or net job changes for women only, or it devises a single ad hoc measure, such as the ratio of male wages to female wages or the female share of exports (e.g. Do, Levchenko, and Raddatz, 2016), even though the motivation for such work relies on how wage changes separately affect income and the cost of time for men and women. Approaches that cannot distinguish male labor demand shocks from female labor demand shocks potentially suffer serious omitted variables biases when the fertility decision depends jointly on male and female income and time.[4] In particular, the relationship between fertility and the woman's wage may depend on whether one conditions on her partner's wage or not, and the same applies to the man.[5]

Furthermore, existing work is not always clear regarding what parameter is being identified, either from a theoretical or policy-based perspective. For instance, increases in wages may induce increases or decreases in hours worked among those already working (especially if the sample comes from a developed country), leading to an ambiguous connection between wages and female time use. Unemployment rates suffer from well-known issues relating to indeterminate changes in the numerator and denominator (since both employment and labor force participation may change across business cycles). Wages or earnings, especially in a setting like Mexico where labor force participation among women remains low, can suffer from severe composition bias (Solon, Barsky, and Parker, 1994). To address these issues, my analysis focuses on changes in net formal sector employment for men and women. Thus, I identify the net impact of increases in formal sector employment for men and women on fertility.

This paper also contributes to a large literature (e.g. Butz and Ward, 1979; MacDonald,

[4] Had I used a measure of the ratio of men's to women's employment, I would have incorrectly found that short-run growth in women's employment decreases fertility, when the effect is driven by short-run growth in men's employment increasing fertility.

[5] Summarizing a major strand of the literature, Jones et al. (2010) note that the correlation between fertility and the wife's wage is usually negative, whether one conditions on the husband's wage or not; the unconditional correlation between the husband's wage and fertility is negative; and lastly, the correlation between the husband's wage and fertility, conditioning on the wife's wage, is either positive or negative, depending on the study.

1983, Macunovich, 1995; Currie and Schwandt, 2014) relating fertility rates to the business cycle, and whether fertility is pro- or countercyclical remains an open question in the literature. Ahn and Mira (2002) note that in their sample of OECD countries, the relationship abruptly switched from being negative to positive. They link the change to a new equilibrium of high unemployment and higher female labor force participation in the OECD. Most of this work focuses on wealthy nations, and one of the contributions of this study is to fill the gap by focusing on a developing country. Moreover, much of this literature focuses on time series regressions that take the country as a unit of observation, but these studies do not identify the impacts of local labor market demand shocks on fertility; for instance, nation-wide changes in policy on employer-provided childcare can induce changes in work and fertility that are unrelated to demand shocks.

Finally, because the time and resource costs of children extend across time[6], it is important to consider how a dynamic theory alters the predictions of the static model. Consider a woman who loses her job or whose hours of work decrease. In the near term, the substitution effect may dominate the income effect and lead her to increase time in child care, potentially increasing fertility. However, if the job loss or wage reduction is due to a widespread negative economic shock, such as a recession, it may lead her to change her expectations about her future employment prospects. If she now anticipates permanently lower income in the future, then the job loss could lead the income effect to dominate the substitution effect and hence lead to lower fertility. A more nuanced alternative is that she may increase fertility in the near term, when the opportunity cost of her time is lower, but decrease fertility in the long-run.[7] The findings in this paper provide support for exactly this interpretation. In sum, changes in economic conditions can also alter expectations about future economic possibilities, which may also affect how families space births across time.[8]

[6] Becker (1960) originally compared children to durable goods. Álvarez-Parra, Brandao-Marques, and Toledo (2013) show that spending on durable goods is particularly volatile in developing countries, including Mexico, which is consistent with my finding that fertility strongly responds to the business cycle.

[7] The presence of liquidity constraints or uncertainty about future outcomes, however, may prevent households from fully optimizing across time, implying that even shocks of a short duration matter for households.

[8] An additional line of thinking, dating back at least to Leibenstein (1957), notes that intergenerational

1.2 Data, background, and trends in employment and fertility

1.2.1 Data description

The fertility data in this paper come from the Mexican National Institute of Statistics and Geography (INEGI). These vital statistics data contain individual-level information on all births, including parents' ages, education, type of union, and the municipality they live in. I use the month of birth variable and lag it by nine months to proxy for the year of conception.

Demographic data on municipalities come from the 2005 and 2010 Mexican censuses, and demographic information for intercensal years is linearly interpolated. Employment data come from two sources. The main analysis uses data from the Mexican Social Security Institute (IMSS), which administers the provision of health care, pensions, and social security. All employees in the formal private sector are obligated to register with the IMSS, so these administrative, job-level data contain employment information on the universe of jobs in the formal private sector. I also use data from the Survey of Occupation and Employment (ENOE), which tracks Mexico's labor force and provides detailed information on the characteristics of employment. Unlike the IMSS data, it has the advantage of providing information on employment in all sectors of the economy, but as a survey it contains only a small proportion of formal sector employment and lacks information on some municipalities entirely. The main analysis focuses on years 2005 to 2013, which is the set of years for which IMSS and fertility data are available.

Data on maquiladora line-employment come from the Maquiladora Export Industry Dataset. INEGI collected these data at the monthly level from all export-assembly plants in Mexico from 1990 to 2006. These data contain plant identifiers[9] as well as a variety of

transfers play a major role in encouraging fertility in developing countries. Although I do not pursue the study of the intergenerational transmission of fertility here, it is worth keeping in mind that in countries like Mexico, where some individuals belong to a formal sector that provides social security, and others belong to an informal sector without similar protections, the implications of increased jobs and changes in wages may differ in regions with few formal sector jobs.

[9] Data with plant identifiers must be accessed on-site at INEGI's microdata lab in Mexico City.

information on inputs, expenditures, sales, and value-added. The period of study encompasses major changes in Mexico's exposure to trade, including the signing of NAFTA, the peso crisis, and China's entry into the WTO.

1.2.2 Background on fertility and employment

Figure 1 shows the tremendous variation in general fertility rates[10] across municipalities across Mexico in 2010. These, unsurprisingly, are highly correlated with local incomes, stocks of human capital, urbanization rates, and proportion of speakers of indigenous languages. Mexico's total fertility rate peaked at above seven children per woman in the early 1960's and then entered a period of sharp, continuous decline; it is currently estimated at being just over two children per woman. General fertility rates are also declining: Figure 2 shows the overall trajectory in general fertility rates for the country since 2006, and Table 1 lists fertility rates for federal entities in 2005 and 2013. The most recent trend continues to be downward, but the overall trend in Figure 2 masks substantial variation across the country. For instance, splitting the sample of municipalities into quartiles by levels of urbanization indicates that mostly rural municipalities actually experienced a positive bump in fertility in 2009, the year of the U.S. financial crisis spilling over into Mexico.

While fertility was steadily trending downward prior to 2009, the formal sector employment to population ratio was increasing for both men and women, as indicated in Figure 3. Formal sector employment then contracted sharply in 2009 and has been slowly recovering since. Manufacturing employment in the formal sector has experienced the same trends, but witnessed a much sharper decline in 2009. As has been the case in the U.S. in recent recessions, the men's employment to population ratio suffered a substantially larger decrease in 2009, though within manufacturing women experienced a slightly sharper decline. The overall trends for men and women in employment, though, are similar, which makes connecting aggregate changes in fertility to changes in either men's or women's labor market

[10] I define the general fertility rate as the number of births per 1000 women of ages 15-44.

opportunities (without conditioning on the other gender's job prospects) problematic.

1.2.3 Employment and fertility correlations

Consider, as a benchmark, the case where fertility is a function of overall job opportunities. I first estimate the following regression:

$$y_{m,t} = \lambda + \beta E_{m,t} + \boldsymbol{\alpha} f(X)_{m,t} + \gamma_m + \delta_t + \theta Trend_s + \varepsilon_{m,t} \tag{1.1}$$

The outcome is the natural logarithm of the general fertility rate[11], measured as the number of births per 1000 women aged 15-44.[12] $E_{m,t}$ measures the natural logarithm of total formal sector employment in municipality m in year t for individuals aged 15-44. I also include municipality fixed effects, which control for time-invariant unobservable characteristics specific to each locality; year fixed effects, which control for annual shocks to fertility; and linear state time trends, which control for smoothly evolving determinants of fertility that vary across states (e.g. if regions with high fertility are converging to the fertility rates of regions with low fertility).[13] Standard errors are clustered at the municipality level to account for serial correlation within municipalities (Bertrand et al., 2004).

I do not calculate an unemployment rate as is commonly done in the literature (e.g. Dehijia and Lleras-Muney, 2004; Örsal and Goldstein, 2010; Schaller, 2016) since I only have administrative data on the formal sector, which accounts for fewer than half of all jobs in Mexico. To account for the potentially complex relationship between formal and informal sector jobs, I flexibly control for the natural logarithm of the population of men and women aged 15-44 in the function $f(X)$. These regressions are weighted by the population of women aged 15-44 in each municipality, averaged across years. I limit the sample to

[11] Results are similar if using the raw fertility rate. Using natural logarithms on both the lefthand and righthand sides facilitates the interpretation of the coefficients of interest as elasticities.

[12] This is equivalent to simply using the log of the number of births when the log of population of women aged 15-44 is used as a control.

[13] Results are unchanged when using quadratic time trends, and results using state-by-year fixed effects, which are shown in the robustness section, are similar.

those municipalities with employment in manufacturing to make results comparable between samples using all employment or only employment in manufacturing.

The results in Column 1 in Table 3 show that fertility is procyclical, and whether changes in population are accounted for linearly or in a more flexible way does not alter the main results. Following the framework established earlier, I examine whether manufacturing jobs have a similar effect on employment and re-estimate equation (1) using only formal sector jobs in manufacturing. The results are qualitatively similar using both measures of employment, although the magnitudes are much smaller when focusing on the subset of the formal sector in manufacturing, which is in line with manufacturing employment making up a smaller fraction of the employment stock in municipalities.[14]

Labor markets in Mexico are segmented by sex.[15] For instance, production in garments, toys, musical instruments, perfumes, and cosmetics are the most female-centric sectors within manufacturing, while alcohol, automotive, and concrete manufacturing are the most male-intensive. As a result, labor market opportunities for men and women differ widely across municipalities and time, depending on the share of industries in each location and the growth rate in employment across time. To test whether men's and women's employment have differential impacts on fertility, I estimate the following specification:

$$y_{m,t} = \beta_{male} E_{male,m,t} + \beta_{female} E_{female,m,t} + \boldsymbol{\alpha} f(X)_{m,t} + \gamma_m + \delta_t + \theta Trend_s + \varepsilon_{m,t} \tag{1.2}$$

Results are in Table 3. The estimates for female employment are negative but small and statistically insignificant, while the estimates for men are larger and statistically significant at the 1% level. Note that the magnitudes can be interpreted as elasticities. Controlling for female employment and changes in population, the results imply that a 10% increase in

[14] Strictly speaking, this interpretation depends on the omitted part of employment being orthogonal to manufacturing, which is unlikely to hold due to spillovers among sectors, but is a rough approximation of effects. In a later section, I show instrumental variables results as well as results using only maquiladoras, which relies on expansions and contractions that are arguably exogenous to shifts in other sectors.

[15] Labor market segmentation by sex is common across all countries, even those with high degrees of male-female equality in other spheres, but developing countries in Latin America are especially likely to show segmentation (World Bank, 2011).

formal sector employment for men is associated with a 0.3% increase in fertility rates in that municipality.

These regressions omit the informal sector, which may have differential impacts on fertility. Employment in the informal sector tends to be countercyclical, and negative shocks to the formal sector historically have not led to significantly higher unemployment rates in Mexico. Instead, the informal sector has served as a "safety valve" for individuals on the margin of losing formal sector employment.[16] Since this implies that changes in employment in the informal sector are themselves outcomes of changes in employment in the formal sector, estimates that include both informal and formal sector employment should be interpreted with caution.

Under the theory outlined here, employment in the informal sector should have a smaller impact than the formal sector on fertility. Even if informal sector jobs do not pay less than the formal sector[17], they do not typically provide childcare, social security, and health benefits, so the expected income effect for men is smaller. Women are less likely to have strong attachments to the labor force in the informal sector, so the expected negative effect (if substitution effects dominate income effects) on fertility from growth in the informal sector is likely to be smaller as well. Table 3 shows results combining the formal sector data in the IMSS and data on the informal sector from the ENOE labor force survey, disaggregated by sex. They are consistent with the theory: estimated coefficients for both male and female employment in the informal sector are close to zero. Furthermore, the inclusion of the informal sector does not alter the main results using employment only in the formal sector.[18]

[16] Early models of labor market segmentation construed the informal market as a separate market that serves as an alternative to individuals who cannot enter the formal sector, but more recent work has emphasized that many workers move voluntarily between the formal and informal sectors (Maloney, 2004). Time series of employment indicate that, perhaps due to the stickiness of wages in the formal sector, employment in the informal sector expands during economic crises (Binelli and Attanasio, 2010).

[17] Marcouiller et al. (1997) discuss wages in the formal and informal sectors in Mexico.

[18] Note the labor force survey does not cover all municipalities in Mexico and only goes back to 2006 in its current version, so the sample of included municipalities and years is smaller.

1.3 Identifying the causal impact of local labor market shocks

1.3.1 A measure of predicted employment

The analysis so far has shown that fertility is positively correlated with employment in the formal sector, and this relationship is driven entirely by male employment. Such a relationship may not be causal, however. Using local aggregate employment, as opposed to own-employment, alleviates a major concern about reverse causality: a parent may change his or her own labor force participation in anticipation of or as a consequence of having a child.

Nevertheless, using aggregate measures of employment may still not identify the causal impact of local labor demand shocks on fertility. Potential sources of bias for the fixed effects estimator include simultaneity, omitted variables, and measurement error. Simultaneity may arise if women decrease childbearing, leading some to enter the labor force differentially across municipalities and time in a manner not fully accounted for by the fixed effects employed here. Such a relationship would lead to a downward bias in estimates on women's labor demand.

Omitted factors correlated with labor demand and fertility may also bias estimates. For instance, La Ferrara, Chong, and Duryea (2012) find that telenovelas decrease fertility in Brazil. This points to the potential importance of social norms in shaping households' decisions about women working outside the home and about fertility preference. Furthermore, linearly interpolating the population between census years may smooth temporary shocks to population. Such shocks are likely to be positively correlated with employment and the number of births, leading to an upward bias in the ordinary least squares estimator for both male and female employment.[19]

Finally, measurement error in the variable measuring formal sector employment can induce biases in estimates of both men's and women's employment. The administrative data

[19] Replacing the census population figure with the population estimate from an average of the four quarters in the subsample of municipalities and years in the ENOE does not alter the results here.

cover the universe of formal sector employment, so in that sense they are measured without error. However, some of the movements into and out of the formal sector may represent firms' selectively registering with the formal sector in some periods and not in others. This likely would lead to attenuation bias in the estimates of both coefficients.

To deal with these sources of bias and isolate the impact of changes in local labor demand, I employ two identification strategies in this paper. The first strategy builds on the approach originally employed by Bartik (1991) and used in Blanchard and Katz (1992), Bound and Holzer (2000), and others in urban and labor economics in constructing an instrumental variable predicting labor demand using a "shift-share index." I define

$$Log\ (predicted\ employment)_{m,g,t} = log\left[\sum_{ind} \frac{Emp_{m,g,ind,t=0}}{Emp_{g,ind,t=0}} (Emp_{g,ind,t} - Emp_{m,g,ind,t})\right]$$

to predict the log of employment for group g (men or women) in each municipality m at each time period t. The numerator in the fraction is equal to employment of group g in municipality m in industry ind at time 0, that is, the baseline period. The denominator in the fraction is equal to national employment of group g in industry ind in the baseline period. If one ignores the last of the two terms in parentheses, then in the baseline period this expression is equal to actual employment. In subsequent years it deviates from actual employment because the mix of industries in each municipality is kept constant to address endogenous changes in the industrial mix resulting from local changes in labor supply. The instrument predicts local employment by weighting national employment in each industry with the proportion of employment in that industry located in the municipality in the first period and summing over all industries, separately for men and women. As is common in the literature, I subtract local employment from national employment in parentheses so that the predicted labor market outcome excludes actual local employment. Otherwise, part of the association between the instrument and actual employment would be mechanical.[20]

[20] In practice, local employment in each municipality is small enough that it makes no difference for the outcomes in this paper whether it is included in the expression or not, although the power of the instrument is of course stronger when own-employment is included.

This methodology relies on a municipality's industrial mix in the baseline period predicting outcomes for local workers in subsequent periods. That is, if one municipality has a large employment share in sectors that employ women, such as textile manufacturing, and the employment of women in textile manufacturing increases across the country, we would expect local employment of women to increase. Assuming workers in one sector are comparable across the country, such that a positive (negative) shock nationwide translates into a positive (negative) shock locally, the IV should predict actual local employment.

To satisfy the exclusion restriction, the predicted employment measure must not be correlated with local labor supply shocks. This requires that national changes in employment in a given industry are not due to changes within a single municipality. Mexico consists of over 2,000 municipalities, with the largest accounting for a little over 1% of the population, so this is much less of a concern here than in similar studies on the U.S. using each state as the local labor market. A more subtle point, made in Cosman (2014), is that localities with a similar mix in sectors may be related in other ways, leading the instrument to pick up differences associated with a particular industrial mix rather than differences in employment in that industry.[21] He concludes that Bartik-style instruments do indeed predict local changes in employment from national shocks to industries.

In principle, the shift-share index can be created for any set of industries, but the argument behind it relies on local changes in employment being sensitive to national trends. Local industries, such as those in services, are less likely to respond to national changes in the same industry than sectors that are traded nationally or globally. A demand shock to internationally traded products, in particular, is likely due to exogenously determined factors that lead to a push in local employment in sectors making those products. Hence, I create two measures of predicted employment: one set for men and women employed in the formal sector, as well as a second set focusing only on manufacturing employment in the formal sector. Focusing on manufacturing provides other advantages. First, different types

[21] Using Monte Carlo trials, Cosman investigates whether such a correlation would lead to a spuriously strong first stage and finds that the coefficient on the instrument would actually be negative.

of manufacturing in Mexico, as in the U.S., are centered on particular regions, leading to spatial variation in how susceptible places are to exogenous demand shocks. Second, some industries within manufacturing have been in decline in Mexico during this period, perhaps due to competition from low-cost Asian producers (especially after China's entry into the WTO), but the shocks have not been felt equally across all industries, leading to another source of spatial variation within regions containing manufacturing employment. Third, to the extent that different types of employment differentially impact fertility decisions (Lim, 2009), the analysis answers a well-defined question: what impact do the expansion and contraction of manufacturing jobs have on fertility? Fourth, focusing on manufacturing makes the results comparable to a growing literature in economics evaluating how the liberalization of trade regimes, leading to large growth in low-skill factory jobs for young women in developing countries, has affected these women's lives.

1.3.2 Results using instrumental variables

Table 4 shows the first stage results, as well as the instrumented results and the reduced-form for equation 1 using all formal sector employment. Predicted employment is highly correlated with actual employment, and the first stage appears strong. The instrumented regressions indicate that the elasticity of fertility with respect to employment is slightly below 0.2. This coefficient is much larger in magnitude than the OLS coefficient, a point to which I return below.

The theory predicts that exogenous shocks to male and female labor demand differentially affect fertility decisions. If employment for men mainly translates into income effects, then we may expect that positive shocks to men's labor demand should increase fertility. Positive shocks to women's labor demand may increase the opportunity cost of having a child, while also leading to positive income effects, so the net impact of changes in women's labor market opportunities is a priori indeterminate. To examine the theory, I return to equation (9) and instrument employment for men and women with predicted employment for each group.

First stage results using full formal sector employment and only manufacturing employment are presented in Table 5. When focusing on all formal sector employment, the IV for women is strongly correlated with actual female employment, while both IVs have some predictive power for male employment, although the IV for male employment is greater in magnitude and slightly more precisely estimated. These results are not surprising: since labor markets are not perfectly segmented, labor market shocks for one gender are correlated with labor market shocks for the other gender. When including only manufacturing employment, only the male IV predicts male employment and only the female IV predicts female employment.

The results of OLS, IV, and reduced form regressions for all formal sector employment and for manufacturing employment only are in Table 6. The IV coefficients indicate that a 1% increase in formal sector employment for men raises fertility rates by about 0.3%, depending on the specification, while an increase in female employment has a small and statistically insignificant impact on fertility rates. The impacts from manufacturing employment are roughly half of those including all formal sector employment. To put these numbers in perspective, Schaller (2016) finds that a 1% increase in unemployment, using a state-level analysis in the U.S., decreases birth rates by 2.6%. When she disaggregates by gender, she finds that decreases in male unemployment raise fertility, whereas decreases in female unemployment lower fertility, with stronger effects for men. My results are broadly consistent with hers, although the use of unemployment rates in her study makes the magnitudes not directly comparable with mine.

1.3.3 Reconciling the IV and OLS estimates

The OLS and IV results presented so far differ markedly in magnitude, which bears investigating. It seems unlikely that a local average treatment effect is inducing such substantial heterogeneity in responses. It is also difficult to think of an omitted variable that is more correlated with the instruments than the raw employment numbers leading to results of such magnitude.

If the Bartik-style construction of predicted employment is weakly correlated with actual employment, then the instrumental variables estimators can be very inconsistent. As documented in Bound et al. (1995), the IV estimates are biased in the direction of OLS in finite samples if the instruments satisfy the exclusion restriction. If the instruments do not fully satisfy the exclusion restriction, the degree and direction of inconsistency depends on the correlation between the instruments and the error term in the "structural" equation. It cannot be tested directly if the composition of industries within a municipality is correlated with an omitted variable that also affects fertility, but it is difficult to argue that the measure of predicted employment used here leads to a more inconsistent estimator than actual employment in establishing demand shocks. In that sense, the reduced form results can be construed as a bound on the results. Moreover, if the instruments do not satisfy the exclusion restriction, then they are likely to be biased in the same direction. Yet the estimated coefficients on male employment become much more positive and the estimated coefficients on female employment become much more negative.

The most plausible explanation for the discrepancy in magnitudes appears to be measurement error in the employment data. Although these data are at the administrative level, they only comprise the formal sector. Firms, especially smaller ones, can move into and out of the formal sector in Mexico. Some component of what may appear to be job gains or losses can simply be the result of how firms choose to classify themselves. It is well-known that panel data methods amplify the effects of measurement error. The higher the correlation between employment levels in adjacent time periods, the more inconsistent the fixed-effects estimator of the effects of employment is likely to be.

To investigate this more formally, consider a stripped-down first-differenced version of equation (9) regressing changes in fertility on changes in log employment:

$$\triangle y_{m,t} = \beta \triangle Emp_{m,t} + \triangle \varepsilon_{m,t}$$

It follows that

$$plim\ \hat{\beta} = \beta - \frac{\beta\sigma_v^2}{(1-\rho)\sigma_{emp}^2+\sigma_v^2}$$

where σ_v^2 is the variance of the measurement error (for a derivation of this type of measurement error, see Cameron and Trivedi, 2009) and ρ is the correlation between employment in adjacent time periods. The correlation between the logarithm of employment in adjacent years is close to one for some pairs of years. The equation above shows that such high serial correlation leads to a strong attenuation bias for the estimated coefficient on employment. A back-of-the-envelope calculation using the fixed effects estimates and the IV estimates in Table 4 and the variance in employment in adjacent time periods in the data (9.24) leads to an estimated variance in measurement error of 0.54. Even though the measurement error described here is not exactly classical, the back-of-the-envelope calculation appears reasonable given the magnitude of the variance in reported employment and how firms are classified, and it is consistent with the instrumental variables estimates increasing by such large magnitudes.

1.3.4 An approach robust to weak instruments

In this section I propose an alternative set of results that is robust to potentially weak instruments. As noted previously, the large difference in magnitudes between the OLS and IV estimators is unlikely to be due to weak instruments. However, the values of the Kleibergen Paap Wald F statistic, especially for the estimates using all formal sector employment separately for men and women, may be a concern. A traditional rule of thumb from Staiger and Stock (1997) rejects the hypothesis of weak instruments if the first stage F statistic is above 10. That rule of thumb was revised in Stock and Yogo (2005), who formalize the arguments in Staiger and Stock (1997) and provide two criteria for establishing the presence of weak instruments. First, an instrument can be construed as weak if the relative bias of the IV estimator exceeds some level (say 10%) of the bias in the OLS estimator (bias test), and second, an instrument can be thought of as weak if the size of the Wald test exceeds a particular threshold (size test). Stock and Yogo suggest a test statistic that is equivalent to the first stage F statistic (if there is one endogenous regressor) or to the Cragg-Donald F statistic (if there is more than one endogenous regressor) and provide critical values for the F statistic based on the number of endogenous regressors and instruments, the maximum bias allowed (if using the bias test), and the estimation method used.

However, the critical values provided by Stock and Yogo crucially depend on the assumption of conditional homoscedasticity. In models containing heteroscedasticity, serial correlation, or clustering, as is the case in this paper, the critical values are no longer valid. Instead, I follow Kleibergen and Paap (2006), who suggest an F statistic that is robust to the presence of non-independent and identically distributed standard errors. To the best of my knowledge, however, the econometrics literature has not generated a formal test for the presence of weak instruments when errors are not i.i.d. and there are multiple endogenous regressors (see Montiel Olea and Pflueger, 2013, for a recent contribution when there is only one endogenous regressor).

An alternative to testing for weak instruments involves the construction of confidence sets that are robust to weak instruments; this approach exploits a duality to hypothesis testing. Given a test of $\beta = \beta_0$, one can create a confidence set for all values of β_0 for which the hypothesis is not rejected. Several tests have been proposed, including the conditional likelihood-ratio test (Moreira, 2003), the Lagrange multiplier test (Kleibergen, 2002; Moreira, 2002), and the Anderson-Rubin test (Anderson and Rubin, 1949). I present results from the latter because, unlike many of the other tests, it is generalizable to the case of more than one endogenous regressor, uses standard F critical values, and is regression-based, making the implementation straightforward.

Analysis using formal sector employment separated by gender has the smallest values of the Kleibergen-Paap Wald F statistic. This is likely due to the non-tradable sector being included in employment measures, as well as collinearity between male and female measures of employment. Thus, I redo the instrumental variables analysis in Table 6 (that is, of equation 9) by inverting the Anderson-Rubin test to create confidence sets for coefficients on men's and women's employment (see Baum et al., 2007).[22] The Anderson-Rubin test jointly tests $\beta = \beta_0$ and the exogeneity of the instruments. That is, rejection would occur if the null hypothesis is false or if the instruments are endogenous. The construction of the confidence set is done via grid testing. Figures 4, 5, and 6 show the resulting confidence sets (i.e. coefficients within the acceptance region), as well as the rejection surface, for the effect of employment on general fertility rates. It makes little difference whether population is controlled for via a linear, quadratic, or cubic polynomial. Consistent with the earlier results assuming identification, positive (negative) shocks to employment for men lead to positive (negative) changes in fertility rates; that is, for any potential effect of female employment, the effect of male employment is positive. On the other hand, both positive and negative effects are consistent with female employment, although the bulk of the confidence set falls within negative parameter values for female employment.

1.4 Timing of births and effects on permanent fertility

1.4.1 Birth order and lagged effects

Fertility choice is a dynamic process. In theory, it is possible for changes in wages and employment probabilities to affect total fertility, the timing of births, or both. I do not follow cohorts of women over time and observe how their fertility responds to exogenous fluctuations in employment opportunities. Instead, my data tie yearly birth records to employment data, so I cannot address directly the question of how long-lasting any effects from yearly variations in employment conditions may be. However, I can exploit additional information in birth certificates to indicate whether couples are substituting across years, and whether differential effects across the life-cycle are present.

In Table 7, I show results on fertility rates that separately consider only first births, second births, and births of third or higher parity.[23] IV results focusing only on manufacturing

[22] The procedure is implemented using the command weakiv in Stata provided by Findlay et al. (2013).
[23] I divide the total number of births in each parity group by the population of women aged 15-44.

employment indicate that only births of third or higher order have a statistically significant response to changes in employment for men. IV results using the full set of formal sector employment show progressively larger impacts for men and women for higher parities. If couples were timing births earlier in response to increased male employment, then we might expect women with no previous births to increase fertility, but both sets of results indicate that women with no previous births are the least responsive to changes in employment. Instead, these results are consistent with exogenous shocks to male employment leading to permanent increases in the size of families.

To further investigate whether the results here indicate changes in timing or total effects on fertility, I augment Equation (9) with a one year, two year, and three year lag in male and female employment. Table 8 focuses on the reduced form results using all formal sector employment, with each column introducing an additional lag.[24] Including lags of male and female employment substantially increases the magnitude of both current-year coefficients, implying that a 10% positive shock to the men's labor demand index increases fertility rates by slightly over 3%. The coefficent on predicted female employment becomes negative and statistically significant from zero once a set of three lags is included, implying a 2% decreases in fertility rates from a 10% positive shock to the women's labor demand index. The net effect for women (the sum of the lags), however, is close to zero and statistically insignificant from zero.

These estimates provide evidence in favor of two points. First, current-period labor market structure has the biggest impact on fertility decisions. This may be due to liquidity constraints that prevent individuals from optimizing across years, to uncertain expectations about future earnings, or simply the salience (in behavioral terms) of current conditions. Second, substitution effects appear strongest in the current period for women, which is in line with theory: a short-run negative shock to employment for women may reduce the

[24] The sample size shrinks with every additional lag. Results are very similar if restricting to a consistent set of years across column. I only show reduced form results, not IV, because all specifications are restricted to contain at most two instruments to avoid multicollinearity problems with multiple instruments. Additional lags are not statistically significant from zero.

opportunity costs of having a child, while a short-run positive shock to employment may induce some women to postpone fertility.

A caveat to these results is that coefficients from finite distributed lag models can be unstable in the presence of serial correlation. Moreover, results from aggregated birth data across years do not capture the same effect as following the same women over time and studying lagged employment shocks for a given cohort. The next subsection turns to the cyclical composition of employment shocks to disentangle effects on the timing of births from effects on total fertility.

1.4.2 Time series properties of employment

It is important to understand the nature of formal sector employment shocks to interpret whether these shocks have permanent or transitory effects. The results in the previous subsection indicate that the largest impacts are on highest-order births, providing evidence in favor of permanent effects. Those results are also consistent with the results in Heckman and Walker (1990), who find that the majority of the effect of men's incomes and women's wages on total fertility in Sweden is driven by the decision to have a third birth.

To gain a better understanding of the time series properties of employment for men and women and explore the effect of employment dynamics, I follow the methodology in Baker, Benjamin, and Stanger (1999), who filter the minimum wage to study how high- and low-frequency cycles in the minimum wage affect the employment-to-population ratio. I decompose the natural log of predicted employment as follows:

$$Emp_{m,g,t} = \frac{1}{2}(Emp_{m,g,t} - Emp_{m,g,t-1}) + \frac{1}{2}(Emp_{m,g,t} + Emp_{m,g,t-1}). \tag{1.3}$$

The first term in parentheses focuses on sharp, high-frequency changes between years to employment, whereas the second term in parentheses, a moving average, emphasizes slower-moving cycles in employment.

When decomposing employment into higher- and lower-frequency components, I find that the overwhelming majority—over 90%—of the variation in the employment measures comes from the lower-frequency component. Moreover, in a simple test of serial correlation (i.e. of an AR(1) process), the coefficient in a regression of predicted employment for both men and women on a lagged term of employment leads to a coefficient of about 0.7, which is statistically significant from zero at the 1% level in both cases. The evidence here, consistent with the discussion in McNown (2003) on time series studies of the relationship between fertility and economic variables, is that employment shocks are highly persistent. An implication of the variance decomposition is that the results in Table 6 are driven by low-frequency employment variation.

To separate the effects of slow-moving versus fast-moving variation in employment, I show results using the filter in Equation (10) for the log of predicted employment in Table 9. As implied by the variance decomposition, the coefficients on predicted employment (Column 6 in Table 6) in the earlier results are close in magnitude to the coefficients on the low-frequency components in men's and women's predicted employment in Table 9. This means that short-run fluctuations in employment are highly correlated over time, and families presumably anticipate that a (negative) positive shock to labor market opportunities today implies a (negative) positive shock to labor market opportunities in the future. This is especially important in the case of fertility decisions, as both the pecuniary and time costs of having children today accrue over a long time horizon.

The results in Table 9 indicate that cycles in employment for men at both high and low frequencies increase fertility rates. For women, the effect is close to zero at the low frequency, but becomes larger, negative, and statistically significant from zero (at the 10% level) at the high frequency. The frequency domain sheds light on which components of employment—that is, slow, secular changes, versus shocks associated with sharp responses to exogenous shifts in demand—drive fertility decisions.

Because men's employment generates income effects for the family, both low- and high-frequency shifts in employment for men should have a positive impact on fertility, and indeed the results are consistent with the theory. When comparing the magnitude of the coefficient on high-frequency men's predicted employment to the coefficient on low-frequency men's predicted employment, I find that I cannot reject the null hypothesis that the coefficients are equal in magnitude. The impact of a low-frequency shock to men's labor market opportunities, which a family may perceive as permanent income shock, may be expected to have a greater impact on fertility than a high-frequency shock, which may appear to a family as transitory. However, families may not be able to time births perfectly, may face liquidity constraints, or may not have perfect expectations about the future. Heckman and Walker (1990) find that increases in men's incomes reduce times to conception and raise total fertility, and it appears that both effects are present here as well.

The decomposition of women's predicted employment into high- and low-frequency components allows us to reconcile the original results, which show no statistically significant impact of increases in labor market opportunities on fertility, with the results of the distributed lag model, which indicate a negative impact in the short run but no meaningful impact on total fertility over a longer time horizon. For women, it appears that fast-moving, unexpected changes induce relatively larger substitution effects. Thus, transitory negative shocks to employment opportunities, as indicated in the distributed lag models and in the high-frequency component of employment, accelerate the the timing of births since the opportunity cost of having children declines during these labor market shocks. Over a longer time period, permanent increases in income roughly offset the increased opportunity costs of having children. These effects on fertility are masked in the earlier regression results, where low-frequency variation dominates the employment measure.

1.5 How do changes in employment affect wages?

The results in this paper indicate that men's employment has a robust, positive impact on fertility, whereas women's employment has weaker, negative impacts. In a wealthy country with high female labor force participation, we might expect that increases in women's employment opportunities generate large income effects that may counteract substitution effects from the increased opportunity cost of time, but this is unlikely to be the case in Mexico. As theory implies that under certain conditions changes in employment may be associated with changes in wages, I directly evaluate the relationship between employment wages, wages, and sectoral gender intensity in this section.

First, changes in relative employment opportunities for women may increase their bargaining power. Research indicates that men have higher fertility preferences than women (Westoff and Bankole, 2002), so if increases in relative labor market opportunities for women lead to shifts in favor of their preferences, we should also expect to see declines in fertility. Although direct changes in wages and employment status can bring about a change in personal bargaining power, women who observe no change in job status or quality (that is, those women who are unaffected on the extensive margin studied in this paper) may still see enhanced bargaining power when their labor market opportunities improve: what matters is their outside option. These inframarginal changes work in the same direction as changes on the margin, so it seems puzzling that results are relatively weak for women's employment.

Do positive shocks to women's employment increase their earnings? To address this question, I evaluate how changes in labor markets in Mexico affect earnings. Since the instruments constructed in this paper exploit compositional differences across sectors, I focus my analysis on differences among sectors. As a starting point, I calculate mean earnings for men and women in each 4-digit sector in the IMSS data and plot the ratio of mean earnings for men to mean earnings for women against the proportion of men in each sector, along with the line of best fit, in Figure 7. The relative size of each circle indicates the number of individuals in each sector. The graph suggests that sectors with relatively more men have less gender-related earnings inequality. I evaluate this claim by regressing the ratio of earnings on the proportion of men in each sector. Table 10 shows that the relationship is strongly negative when pooling all years and comparing across sectors (column 1).

Some authors have argued that sectors that employ mainly women pay lower wages.[25] When I separately estimate how the proportion of men in each sector is correlated with earnings for men and women, however, I find that women's earnings have a small, albeit positive, correlation with how male-dominated the sector is (column 2). It does not seem to be the case that women working in female-centric industries are earning substantially lower earnings. Men's earnings, on the other hand, are strongly negatively correlated with how many men there are in the sector (column 3).

Suppose, for simplicity, that there are two types of jobs within each sector: a high-paying, high-skill type and a low-paying, low-skill type. If men work in both types of jobs, while women are only employed in the latter, then sectors with few men should have disproportionately more men in the high-paying sectors, leading to the type of wage inequality observed in the IMSS data. Alternately, it is possible that both women and men are distributed in high-skill jobs, but women earn lower relative wages in higher-paying occupations. To investigate these possibilities, I turn to the ENOE, a labor force survey in Mexico similar to the CPS in the U.S., to study the relationship between wages, gender composition of industries, and skill intensity.[26] Columns 4-6 in Table 10 reproduce the same results as in the IMSS data: wage inequality is negatively correlated with the proportion of men in each sector, and this is mainly due to men earning less on average in sectors with more men.

Table 11 shows results for regressions of either wage inequality, log of male wages, or log of female wages in each sector against the proportion of men in each sector and the proportion of each gender in high-skill occupations. I define individuals as being in high-skill occupations if they are employed as professionals or managers, which are the two highest paid occupational classifications.[27] Once the proportion of men and women in high-paying

[25] For an extensive discussion of forces shaping gender differences in employment in developing countries, see the World Bank Development Report (2011).

[26] The ENOE contains a different classification of industries from the IMSS and consists of a smaller sample, so the results are not directly comparable. I limit the analysis to those sectors with at least 10 individuals of each gender in all years. Varying the cell size limit slightly does not affect the results qualitatively.

[27] The other choices in the survey are as follows: educational workers, clerks, industrial workers, merchants, transport operators, workers in personal services, protection and surveillance workers, and farmworkers.

jobs is accounted for, the degree of male bias in each sector's workforce ceases to predict either wage inequality or wages for men or women.

To exploit differences across time and space in male versus female specialization, I return to equation (9) and use the IMSS data to evaluate how employment for men and women affects their earnings. Table 12 shows the impact of raw employment itself and the reduced form. Earnings for both men and women are positively correlated with expansions in male employment and negatively correlated with expansions in female employment. The predicted demand measure, which isolates purely demand factors, shows even larger impacts for women, with an elasticity of 0.19 for men's demand shocks and -0.21 for women's demand shocks. Men's earnings are unaffected by changes in demand for male or female labor. If all jobs were identical, a positive shock to demand would be expected to raise wages, but because women occupy lower wage positions, positive shocks to their demand actually lower their wages, conditional on men's demand shocks. In other words, separate demand shocks for women and men likely lead to compositional changes in the labor force. For instance, if low-skill women are induced into the labor force through positive shocks while high-skill women's labor market opportunities remain unaffected, then average female wages may decline, even though wages for all women have weakly increased.

The difference in results for men's wages when focusing on supply and demand shocks versus demand shocks only can be reconciled if male supply is increasing more among educated groups, yet demand is more pronounced in lower-wage positions. This is consistent with the results in Campos-Vázquez (2013), who evaluates trends in wage inequality in Mexico following the passage of NAFTA. He finds that the supply of college-educated workers has grown rapidly, but high-skill occupations have not expanded enough to fill the new demand for these positions, leading to wage compression at the top.

Even if positive demand shocks for men do not translate into higher wages *among formal sector jobs*, formal sector jobs provide substantial health, social security, and childcare benefits that are not available in the informal sector. Second, increased labor force participation on the extensive margin translates into pure income effects. Thus, regardless of the effect on wages within the formal sector, increases (decreases) in male employment should lead to higher (lower) fertility, which is consistent with the results in this paper.

1.6 The impact of maquiladora employment on fertility

1.6.1 Context behind the expansion of maquiladoras

The previous section established how changes in men's and women's formal employment impact fertility decisions. In this section, I describe the history and context behind a particular type of formal sector employment in Mexico's export-assembly plants, introduce a different dataset and estimation strategy, and show results consistent with the earlier set of findings.

After the termination of an agreement in which Mexico sent seasonal farm laborers to the U.S., Mexico faced the prospect of a large pool of unemployed people living in the north.[28] To generate incentives for firms to locate in the northern border region, in 1965 the Mexican government introduced a plan called the Border Industrialization Program, which allowed full foreign ownership of establishments in Mexico. Although these establishments, called maquiladoras, were highly regulated initially, over time these regulations were relaxed to attract more foreign investment.[29]

In particular, Mexico's poor economic performance in the 1980s led to a series of economic reforms and the liberalization of trade conditions, which led to large-scale growth in maquiladora employment starting with the late 1980s. Figure 9 shows the growth of employment for each gender over the time period studied. Although aggregate employment for men and women closely tracked each other, the graph masks how gender-segmented facto-

[28] This was known as the Bracero Program. The combination of demand for low-wage agricultural workers in the U.S and excess supply of labor in Mexico allowed the program to operate until 1964, when pressure from U.S. labor unions led to the termination of the program.

[29] For instance, maquiladoras were required to be within twenty kilometers of the border and all output had to be exported. The bulk of employment has stayed near the border, however, as shown in Figure 8.

ries are: for instance, factories specializing in textiles employ mainly women, while factories manufacturing electronics employ mainly men. This segmentation applies not only at the establishment-level, but also at the regional level, as shown in Figure 10, which illustrates the average female share of maquiladora employment across Mexican municipalities.

1.6.2 How maquiladoras affect fertility

The growth in maquiladora line-employment happened at both the extensive and intensive level: new plants opened, and plants that continued to operate increased in size. Figure 11 shows the density of employment at the establishment-level in 1990, 2000, and 2006. Although there is a clear clear shift toward bigger sizes over time, typical sizes of maquiladoras remained small, with median employment below 100 people.

To study the impact of the rapid growth in employment for both men and women across sectors and regions on fertility decisions, I focus on establishment-level changes in employment from 1990-2006. I instrument for net new jobs for women (men) in export assembly plants within each municipality with net new jobs for women (men) in these plants that are solely due to large single-firm expansions/openings and contractions/closings (i.e. a change of least 50 individuals in a given year). As Figure 11 indicates, maquiladoras are quite small, so these are large changes relative to the size of the establishment.[30]

For the exclusion restriction to hold, I require that firms do not respond to fertility decisions (or to any omitted determinant of fertility) with large expansions or contractions, conditional on the fixed effects and controls for population in the estimating equation. This seems especially plausible in my context, as I focus only on maquiladoras. It is likely that large changes in employment at these plants involve high fixed costs and are due to shocks in external demand from other countries (Atkin, 2012).

I create the employment variables as follows. I aggregate male and female employment separately across all establishments within a municipality to create the main independent

[30] Similar identification approaches have been employed by Atkin (2012) and Ananat et al. (2011).

variable for each gender.[31] To standardize the variable, I divide employment by the 1990 population of working-age men or women.[32] I use the 1990 baseline year in the data for two reasons: the main one is that the denominator may vary along with other conditions in the municipality, and since maquiladora employment is very small relative to all employment in most places, this can lead to a severe bias in the variable. (For instance, if maquiladora employment increases for men, but population shifts lead to a large enough increase in the denominator, then the term may decrease even when maquiladora employment goes up.) Second, I interpolate the population between census years, which may introduce an additional source of bias. Hence, using the 1990 population creates a standardized measure to track how changes in maquiladora employment for each gender relative to the baseline population of that gender affect fertility and marriage outcomes.

To construct the instrumental variable, I first difference employment for each gender at each establishment across years. Keeping only the sample of establishments that contains a change of at least 50 individuals from the previous year, I then aggregate employment to the municipality level. Finally, I standardize by the 1990 population of that group, as with the main independent variable.

Table 13 shows the first stage results for the full sample of municipalities in columns 1 and 2. Since a large proportion of municipalities has only one or two small factories, which are unlikely to have a large impact on fertility if the population is large, I also limit the sample to those municipalities that have at least 1% employment of either gender in at least one year, which I call the restricted sample.[33] The results for this sample are in columns 3 and 4. In all cases, the Kleibergen-Paap Wald F statistic is above 50, and both the instruments

[31] I calculate average yearly employment from the monthly observations. Establishments can enter or exit the data in any month, and some establishments do not contain data for some months between entry and exit (i.e. due to temporary shut-downs or not answering the survey, which was required). For the set of plants missing months, I tried three techniques: calculating average yearly employment only for those months in the data, imputing employment based on previous months, or simply imputing zero. Results are similar for all approaches.

[32] The denominator is the population of men or women aged 15-44 in that municipality, linearly interpolated from the decennial census.

[33] Results are similar if restricting to 3% or 5% employment, although the sample size is substantially reduced.

for male and female employment are strongly correlated with actual employment.

Table 14 shows the OLS, IV, and reduced form results for the full sample, as well as the restricted sample. The reduced form results indicate that men's employment has a large, positive impact on fertility, while the impact of women's employment is negative, albeit imprecisely estimated. The results imply that if, say, expansions or openings of new maquiladoras lead to an additional 5% of the population to work in maquiladoras, this would translate into a 0.01 increase in the log of the fertility rate. Although this may seem small, it is important to keep in mind that maquiladoras make up a small proportion of the labor force in most municipalities.

1.7 Migration, robustness, and alternative specifications

1.7.1 Migration

If individuals migrate into or out of municipalities experiencing changes in labor markets, then resulting changes in fertility may be due to changes in the composition of individuals living there and not to actual changes in behavior. There are three types of migration that may affect the results. The first type concerns local migration: individuals may live in one municipality and work in another. Since I link aggregate employment to aggregate fertility in the same municipality, we can think of some births in the data as being assigned to the wrong municipality. If men and women are equally likely to commute to other municipalities for work, this can lead to measurement error that should bias the estimators for both groups' employment toward zero. If men are more likely to commute to different municipalities, as seems likely the case in this setting (unfortunately, I cannot test this directly), then the male estimator should be more attenuated. Since I find larger results for men rather than women, it appears unlikely that this type of bias is driving the results.

A second type of migration concerns moves across the country. As Mexico has industrialized, individuals living in poor rural communities in the south have moved to the north to work in maquiladoras and related enterprises. Suppose some young women migrate to the north to work in factories and then return home to their rural communities, which have little formal sector employment, to have children. This should bias me in favor of finding a larger negative impact of female employment, yet I find no statistically significant impact.

Another possibility concerns selective migration into municipalities that undergo demand shocks. This type of migration can result from either population movements within Mexico, or cross-country migration (such as Mexican migrants returning from the U.S.). To probe this further, I re-run equation (9), except I exclude population controls from the right-hand side and instead use them as the outcome variables. Instrumenting employment for men and women with predicted employment, I find that labor demand shocks for men and women are not associated with changes in population for either men or women (Table 15). Although it is possible that in-migration of a selected sample is exactly balanced by out-migration, these results suggest that migrants are not choosing municipalities based on labor demand shocks, at least in the short run. This is reassuring: changes in raw employment are due to shifts in supply and demand, whereas the Bartik-style instrument should be isolating only changes in labor demand.

1.7.2 Robustness

This paper uses an instrument in the tradition of Bartik (1991) to isolate an exogenous predictor of employment. To test whether the results hold using a traditional Bartik-style instrumental variable, I create

$$Bartik\ instrument = \sum_{ind} \frac{Emp_{m,g,ind,t=0}}{Emp_{m,g,t=0}}\ log(Emp_{g,ind,t} - Emp_{m,g,ind,t}).$$

The definition of the terms is the same as before, except the denominator in the fraction is now equal to employment in municipality m for group g (men or women) in the baseline time period. This instrument is often constructed in first-differenced form and used to predict employment growth. Two municipalities with exactly the same industrial composition would have the same values of this measure (ignoring the subtraction of own-employment in parentheses), as they would be expected to have similar levels of growth or decline in employment; in other words, it is invariant to the municipality's population. However, including municipal fixed effects means the Bartik instrument effectively isolates the same type of variation as the instrument used in the paper, and results replicating Table 6 (shown in Table 16) using the instrument are qualitatively similar, though formal sector estimates are slightly larger in magnitude.

Finally, I replace linear state trends with state-by-year fixed effects and again reproduce the main analysis in Table 17. Such a specification flexibly controls for any unobservable shocks specific to states in any particular year. Of course, municipalities located close to each other are more likely to have similar labor market structure and thus to face similar types of labor market shocks. There is no reason to believe that identification based on the remaining variation across municipalities leads to more consistent estimators, given that state-by-year fixed effects may absorb "too much" labor market variation.[34] In practice, the choice of linear, quadratic, or state-by-year fixed effects matters little: for instance, if including year fixed effects, municipality fixed effects, a cubic in log population for men and women, and year-by-state fixed effects, the impact of male formal sector employment decreases slightly from 0.32 to 0.30, and standard errors increase as well, but the results for men remain statistically significant at the 5% level, and results for women remain small and statistically insignificant. It does not appear that nonlinear, unobservable deviations correlated with labor market changes within states are driving the earlier results using all formal sector employment. Specifications using only manufacturing employment result in very similar estimated impacts for job opportunities for men, but standard errors increase and the coefficients are statistically significant only at the 10% level.

[34] An analogous point is made in Bound and Solon (1999), who show that identifying the returns to schooling based on twin-comparisons may not lead to more consistent estimation.

1.8 Conclusion

The question of how labor market opportunities shape decisions about the family has long interested economists. In some developed countries, fertility rates are arguably too low, and policymakers have invested large sums in relaxing perceived constraints to having children, such as providing flexible working arrangements for young mothers, daycare, or simply lump sum payments. On the other hand, in many developing countries fertility rates remain stubbornly high, making countries that have gone through a large transformation in family structure potentially useful guides for their own experiences.

Mexico has experienced a dramatic fall in fertility, as well as a steady increase in labor market opportunities for young women, driven in part by an expansion in trade-oriented manufacturing jobs. In more recent years, sectors traditionally employing young women, such as textile manufacturing, have become less competitive as production has shifted to China and other Asian economies with lower labor costs. This paper evaluates how expansions and contractions in employment that vary across municipalities in Mexico and over time affect fertility. Because fertility and employment are joint household decisions, I focus on aggregate changes in demand conditions for both men and women, using a measure of predicted employment that exploits labor market segmentation by sex and the industrial structure of each local labor market, to isolate exogenous demand shocks for each group. The findings are robust to an alternative identification strategy based on the expansion and contraction of maquiladoras during the 1990s and early 2000s.

That women's employment does not appear to have a significant net effect on fertility may indicate that other broader, long-run social changes play an even greater role in explaining variation in fertility rates across time and space. In particular, economists and demographers have documented that increases in educational attainment or health (leading to a preference for child quality over quantity), urbanization, and better access to and knowledge of contraceptives have all been associated with declines in fertility.[35] By focusing

[35] For a discussion of theories of the fertility transition, see Mason (1997) and Guinnane (2011) for the

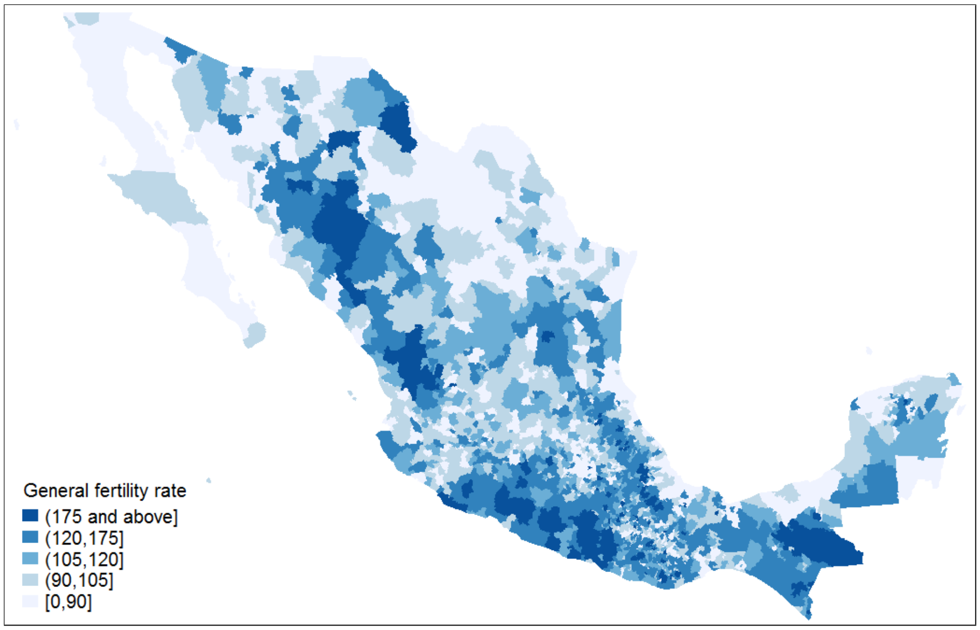
on annual fluctuations in employment within a single country over a relatively short period, which arguably keeps these longer-horizon variables fixed, I am able to identify the causal impact of shocks to labor market opportunities on fertility.

The findings in this paper show that employment dynamics for both men and women matter: positive demand shocks to men's labor lead to positive changes in fertility, and a variance decomposition of employment into high- and low-frequency components provides evidence that increases in men's labor market opportunities in the formal sector lead to higher total fertility. Transitory positive shocks to women's labor, on the other hand, increase the opportunity cost of having children and hence lead families to delay fertility. Evidence from the low-frequency component in exogenous employment variation for women, as well as distributed lag models, indicate that over a longer period the net impact on total fertility is negligible.

perspectives of a demographer and an economist, respectively, and the references therein.

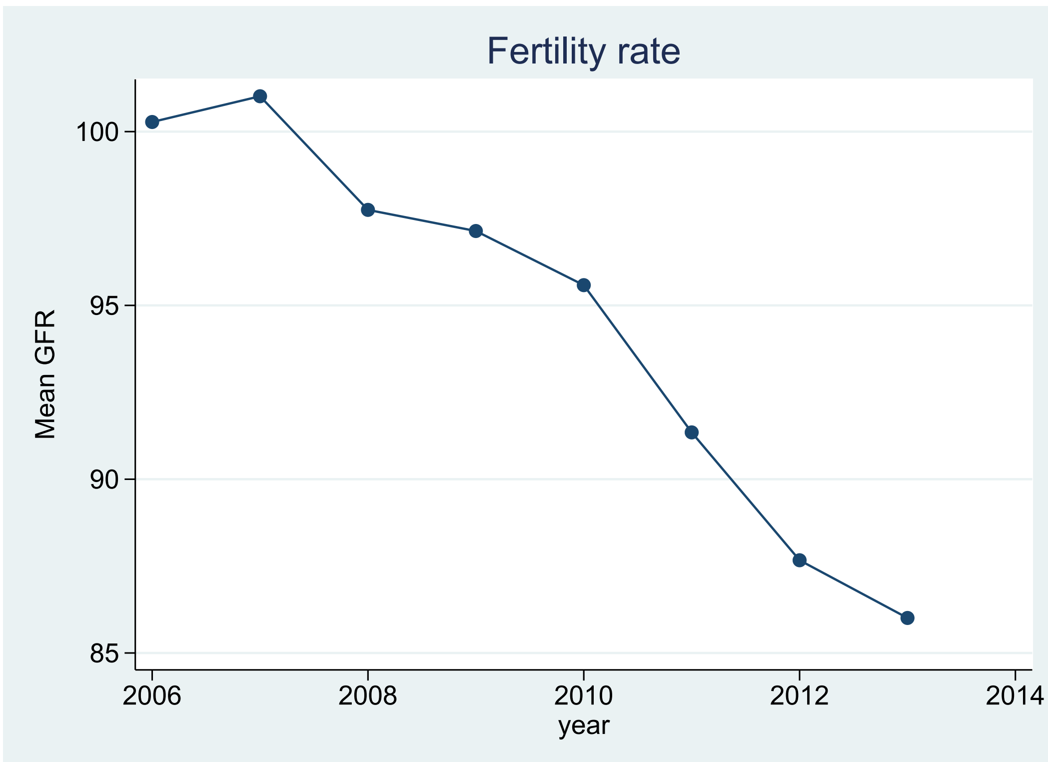
1.9 Figures

Figure 1.1: Fertility rates across Mexico



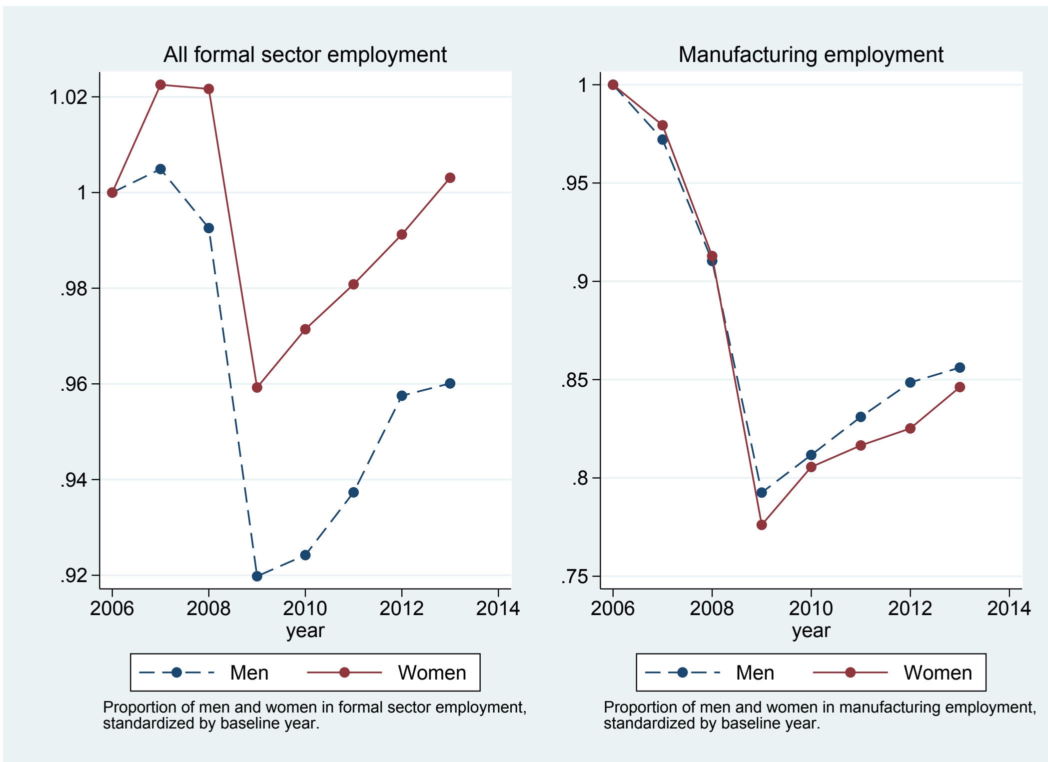
Notes: The general fertility rate is calculated as the number of births per 1000 women aged 15-44. Data are calculated from Mexico's National Institute of Statistics and Geography natality and census statistics in 2010.

Figure 1.2: Fertility rate over time



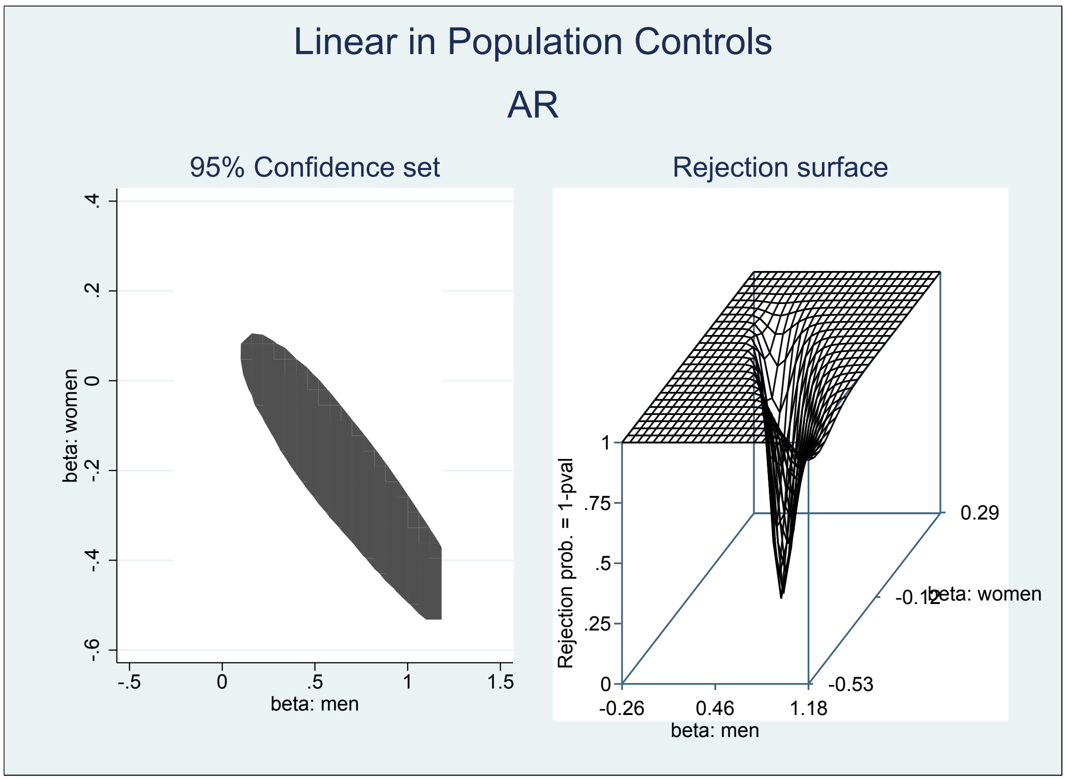
Notes: The general fertility rate is calculated as the number of births per 1000 women aged 15-44. Data are calculated from Mexico's National Institute of Statistics and Geography natality and census statistics.

Figure 1.3: Standardized proportion of men and women in employment



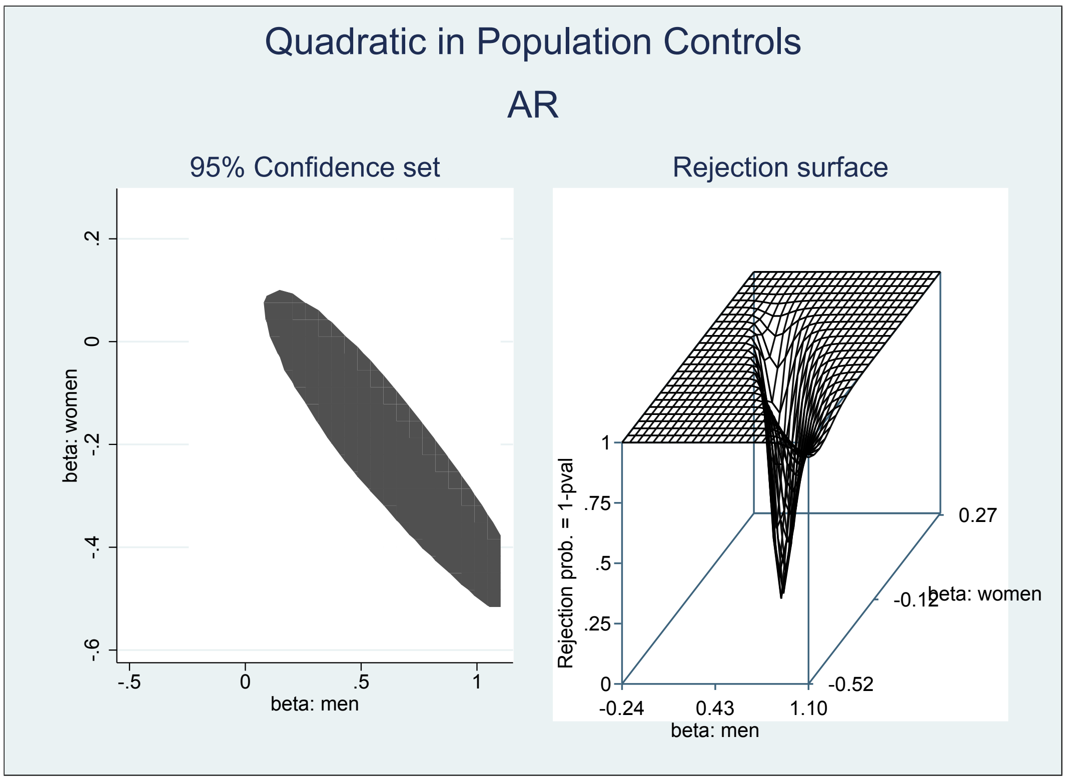
Notes: Trends in formal sector employment, standardized by employment in the first year, are shown. Employment data come from the Mexican Social Security Institute.

Figure 1.4: Confidence set with linear population controls



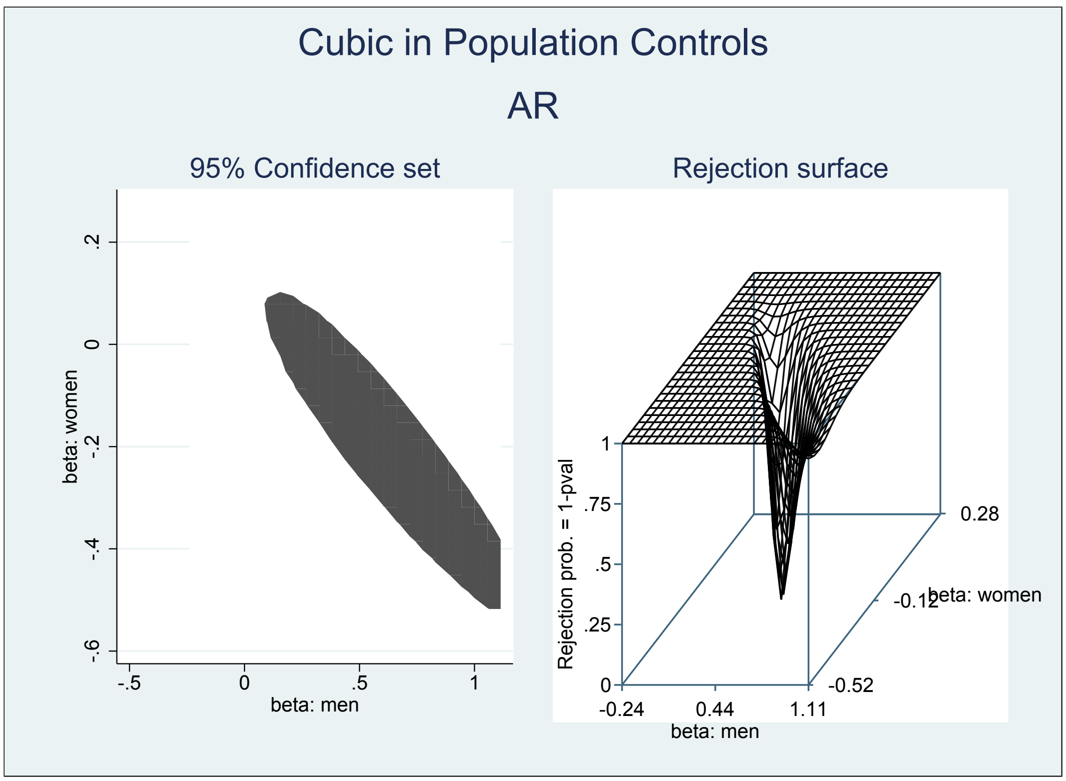
Notes: Confidence set and rejection surface for the Anderson-Rubin test.

Figure 1.5: Confidence set with quadratic population controls



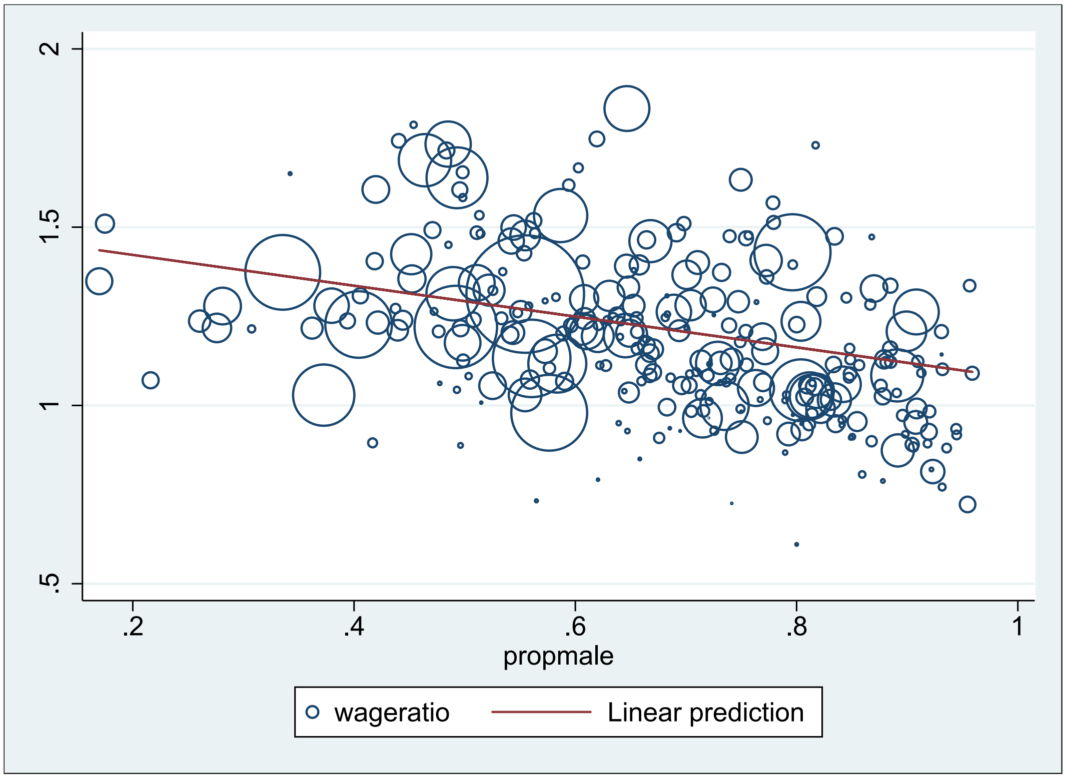
Notes: Confidence set and rejection surface for the Anderson-Rubin test.

Figure 1.6: Confidence set with cubic population controls



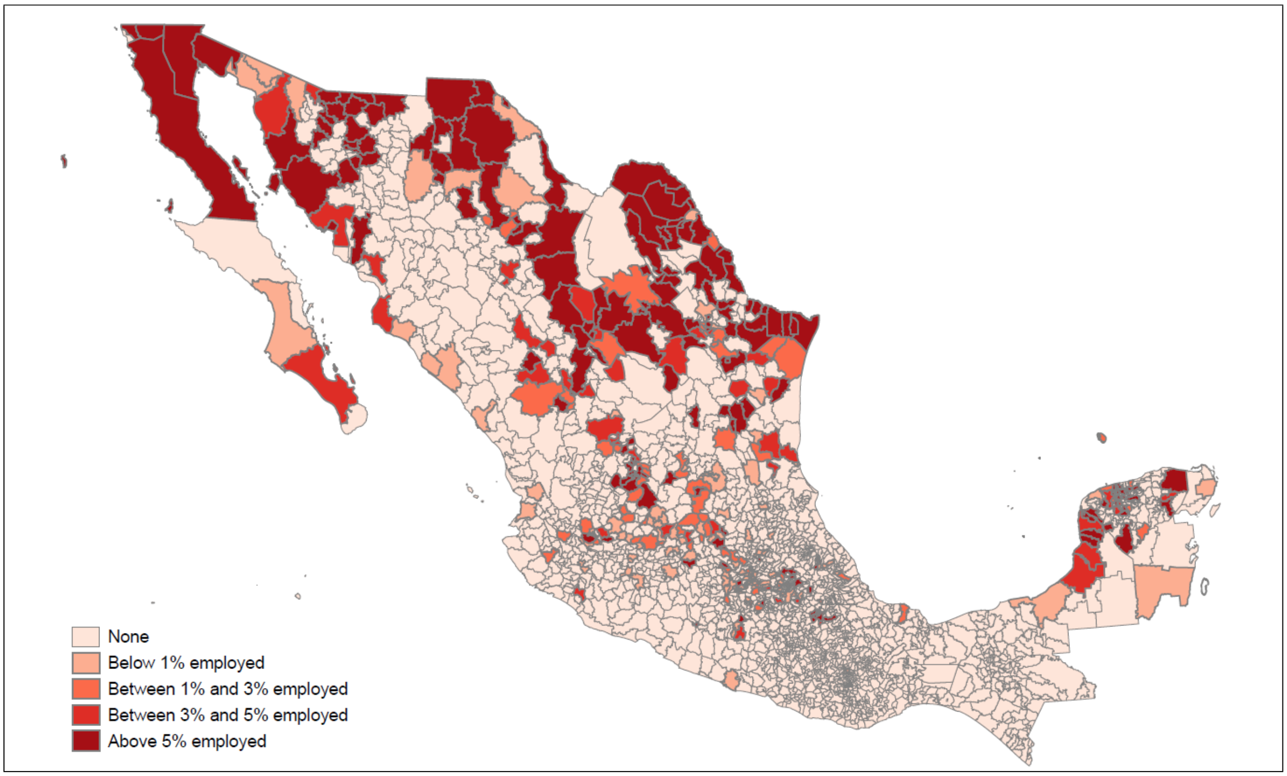
Notes: Confidence set and rejection surface for the Anderson-Rubin test.

Figure 1.7: Male to female earnings ratio in 2005



Notes: The size of each circle represents the size of each sector. The line of best fit is shown. Data are calculated from the Mexican Social Security Institute.

Figure 1.8: Municipalities by maximal share of employment in maquiladoras

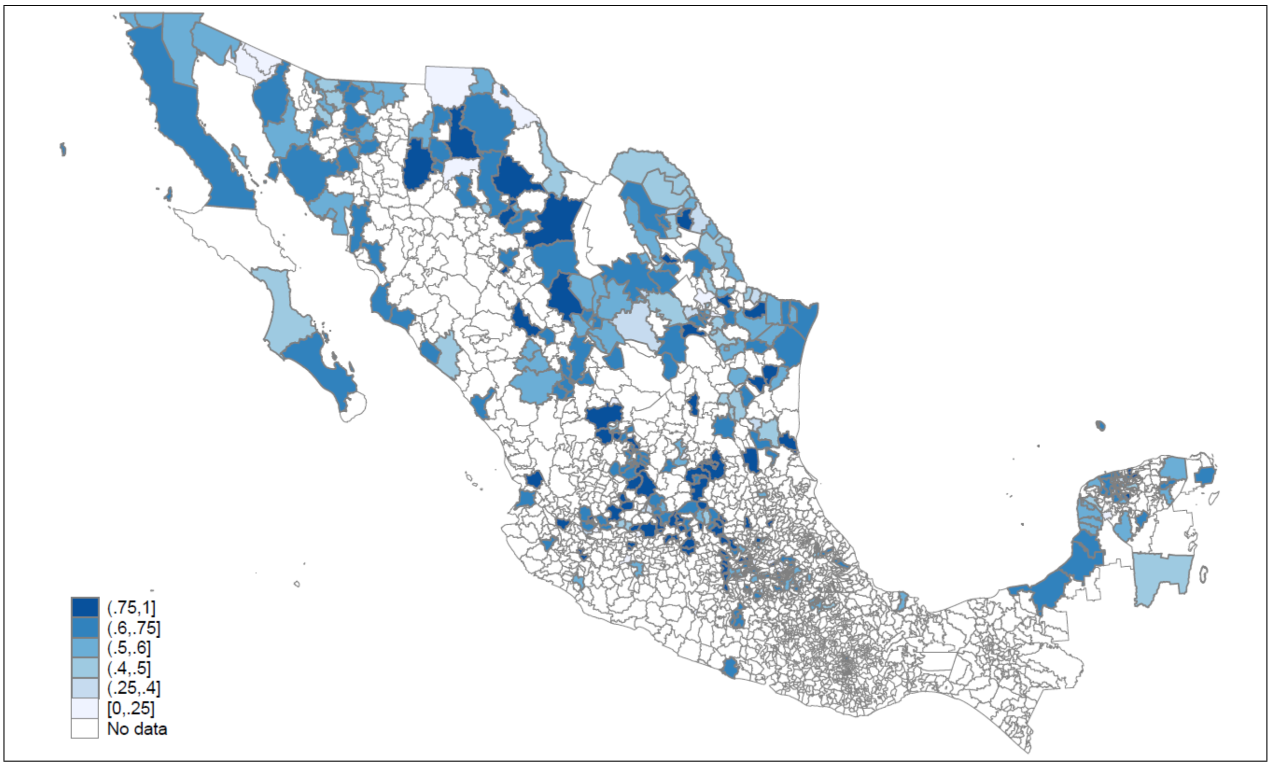


Notes: The share is calculated separately for men and women by taking aggregate employment in maquiladoras and dividing by the estimated population of that group aged 15-44 in each year. Then the maximal share is calculated across all years in the dataset. Data are calculated from the Maquiladora Export Industry Dataset, which is provided by Mexico's National Institute of Statistics and Geography.

Figure 1.9: Employment for men and women in maquiladoras

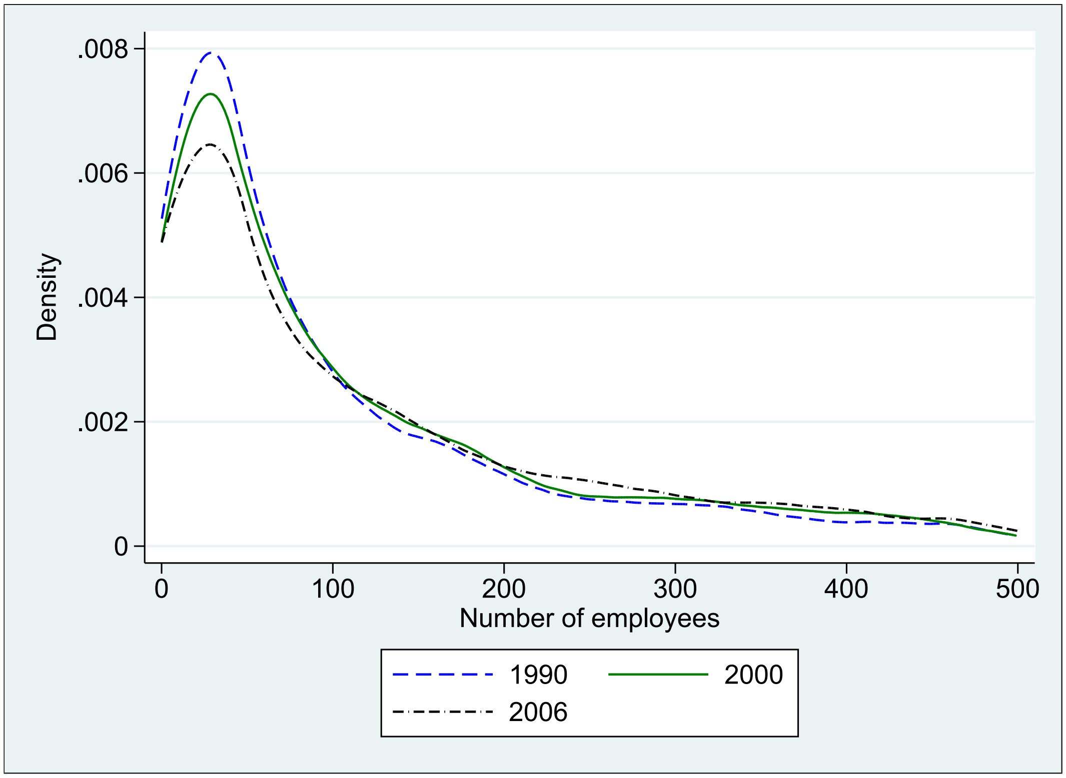
Trends in aggregate line employment for men and women at maquiladoras are shown. Data are calculated from the Maquiladora Export Industry Dataset, which is provided by Mexico's National Institute of Statistics and Geography.

Figure 1.10: Municipalities by female share of maquiladora labor force



Notes: The share is calculated separately for each municipality. The "No data" category refers to municipalities with no maquiladora exmployment. Data are calculated from the Maquiladora Export Industry Dataset, which is provided by Mexico's National Institute of Statistics and Geography.

Figure 1.11: Density of employment across maquiladoras



Notes: The density of line employment at maquiladoras is shown for 1990, 2000, and 2006, using the Epanechnikov kernel with the "optimal width" that minimizes mean integrated squared error if the data were Gaussian.

1.10 Tables

Table 1.1: General fertility rates across Mexican federal entities

State	Fertility rate in 2005	Fertility Rate in 2013
Aguascalientes	93	88
Baja California	89	77
Baja California Sur	94	71
Campeche	87	86
Coahuila de Zaragoza	92	91
Colima	82	78
Chiapas	137	131
Chihuahua	94	79
Distrito Federal	69	66
Durango	108	101
Guanajuato	97	80
Guerrero	149	114
Hidalgo	108	78
Jalisco	91	83
Mexico	92	79
Michoacán	103	94
Morelos	93	82
Nayarit	103	86
Nuevo León	83	79
Oaxaca	112	83
Puebla	115	98
Querétaro	96	81
Quintana Roo	91	76
San Luis Potosí	101	83
Sinaloa	94	83
Sonora	93	81
Tabasco	102	84
Tamaulipas	97	74
Tlaxcala	103	80
Veracruz	102	83
Yucatán	81	74
Zacatecas	101	90

General fertility rates are calculated as the sum of all births per 1000 women of ages 15-44 in each federal entity. The denominator is linearly interpolated from decennial census data from INEGI.

Table 1.2: Summary statistics

	Mean	SD
General fertility rate	89.59	18.45
Formal sector employment	59645	84827
Male formal sector employment	23309	34252
Female formal sector employment	36336	50799
Male manufacturing employment	12332	19789
Female manufacturing employment	7139	13634
Female earnings	171.62	50.68
Male earnings	212.32	62.82
Female population (15-44)	115411	118753
Male population (15-44)	108985	113540
GFR - 1st births only	39.57	9.45
GFR - 2nd births only	24.97	4.87
GFR - 3rd births and above	25.03	10.38
Observations	9,108	
Municipalities	1,012	

Summary statistics are shown across years 2005-2013, weighted by population of women aged 15-44 in each municipality. Author's calculations from Mexico's National Institute of Statistics and Geography data.

Table 1.3: Estimation of the relationship between employment and fertility

	(1) GFR	(2) GFR	(3) GFR	(4) GFR	(5) GFR	(6) GFR
All emp.	0.031***					
	(0.0096)					
Manuf. emp.		0.012**				
		(0.0058)				
Male emp.			0.033***		0.032***	
			(0.014)		(0.014)	
Fem. emp.			-0.0096		-0.00048	
			(0.013)		(0.016)	
Male manuf. emp.				0.022**		0.022**
				(0.0088)		(0.0099)
Fem. manuf. emp.				-0.0055		-0.0037
				(0.0063)		(0.0073)
Male inf. emp.					-0.00073	-0.0012
					(0.0047)	(0.0047)
Fem. inf. emp.					-0.0045	-0.0045
					(0.0039)	(0.0039)
Observations	9,108	9,108	9,108	9,108	5,630	5,630
Municipalities	1,012	1,012	1,012	1,012	811	811
State trend	Yes	Yes	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes

Standard errors are clustered at the municipality level and reported in parentheses. The dependent variable is the log of the fertility rate [number of children per 1000 women of childbearing age]. Main independent variables are formal sector employment, male and female formal sector employment, manufacturing employment, male and female manufacturing employment, and male and female informal sector employment (all in logs). Regressions are weighted by population of women aged 15-44, and results are for years 2005-2013. Cubic polynomials of the log population of men and women aged 15-44 are included as controls.

Table 1.4: First stage and OLS, IV, and RF of impact of employment on fertility

	(1) 1st stage	(2) OLS	(3) IV	(4) RF
Log (pred. employment)	0.85***			0.13***
	(0.091)			(0.033)
Log (all employment)		0.031***	0.16***	
		(0.010)	(0.042)	
Observations	9,108	9,108	9,108	9,108
Municipalities	1,012	1,012	1,012	1,012
State trend	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Kleibergen-Paap Wald F statistic			87.5	

The first stage for the log of employment is shown in the first column. OLS, instrumented results (using log of employment as IV), and the reduced form are shown in remaining columns, where the dependent variable is the log of the fertility rate. Standard errors are clustered at the municipality level and reported in parentheses. The dependent variable is the log of formal sector employment. Regressions are weighted by population of women aged 15-44, and results are for years 2005-2013. Cubic polynomials of the log population of men and women aged 15-44 are included as controls.

Table 1.5: First stage regressions for men's and women's employment

	(1)	(2)	(3)	(4)
			First stages	
	Men	Women	Men (manuf)	Women (manuf)
Log (pred. male emp.)	0.45**	-0.13		
	(0.14)	(0.11)		
Log (pred. female emp.)	0.30**	1.14***		
	(0.10)	(0.11)		
Log (pred. manuf male emp.)			0.54***	-0.11
			(0.15)	(0.18)
Log (pred. manuf female emp.)			0.21	1.18***
			(0.13)	(0.16)
Observations	9,108	9,108	9,108	9,108
Municipalities	1,012	1,012	1,012	1,012
State trend	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes

First stage regressions for the log of male and female employment (columns 1 and 2) and the log of male and female manufacturing employment (columns 3 and 4) using the predicted measure of employment described in the text are shown. Standard errors are clustered at the municipality level and reported in parentheses. Regressions are weighted by population of women aged 15-44, and results are for years 2005-2013. Cubic polynomials of the log population of men and women aged 15-44 are included as controls.

Table 1.6: Impact of male and female employment on fertility

	(1) OLS	(2) IV	(3) RF	(4) OLS	(5) IV	(6) RF
Men (all)	0.0328**	0.32**				
	(0.0139)	(0.13)				
Women (all)	-0.00105	-0.087				
	(0.0129)	(0.079)				
Predicted men (all)			0.15**			
			(0.061)			
Predicted women (all)			-0.0055			
			(0.053)			
Men (manuf)				0.0189**	0.20**	
				(0.00854)	(0.089)	
Women (manuf)				-0.00522	-0.060	
				(0.00606)	(0.052)	
Pr. men (manuf)						0.12**
						(0.047)
Pr. women (manuf)						-0.027
						(0.043)
Observations	9,108	9,108	9,108	9,108	9,108	9,108
Municipalities	1,012	1,012	1,012	1,012	1,012	1,012
State trend	Yes	Yes	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Kleibergen-Paap Wald F statistic		10.0			13.7	

OLS, IV, and reduced form regressions shown. The first three columns include all formal sector employment in logs, while the next set of columns only includes formal sector manufacturing employment in logs. The dependent variable in all cases is the log of the general fertility rate. Standard errors are clustered at the municipality level and reported in parentheses. Regressions are weighted by population of women aged 15-44, and results are for years 2005-2013. Cubic polynomials of the log population of men and women aged 15-44 are included as controls.

Table 1.7: Impacts on fertility rates by birth order

	(1) OLS	(2) IV	(3) OLS	(4) IV	(5) OLS	(6) IV
	1st birth	1st birth	2nd birth	2nd birth	3rd+ birth	3rd+ birth
Men	0.014	0.059	0.073***	0.62***	0.055*	0.83***
	(0.026)	(0.20)	(0.025)	(0.22)	(0.028)	(0.27)
Women	0.027	0.044	-0.021	-0.31**	-0.034	-0.35**
	(0.022)	(0.12)	(0.021)	(0.12)	(0.026)	(0.15)
Observations	9,108	9,108	9,108	9,108	9,105	9,105
Municipalities	1,012	1,012	1,012	1,012	1,012	1,012
State trend	Yes	Yes	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
KP Wald F stat		10.0		10.0		10.0

OLS, IV, and reduced form regressions of fertility rates by parity on the log of predicted employment for men and women are shown. Standard errors are clustered at the municipality level and reported in parentheses. Regressions are weighted by population of women aged 15-44, and results are for years 2005-2013. Cubic polynomials of the log population of men and women aged 15-44 are included as controls.

Table 1.8: Impact of current and lagged predicted employment on fertility

	(1) GFR	(2) GFR	(3) GFR	(4) GFR
Pred. male employment	0.15**	0.32***	0.32***	0.39***
	(0.061)	(0.098)	(0.10)	(0.11)
Pred. female employment	-0.0055	-0.17	-0.15	-0.20**
	(0.053)	(0.11)	(0.096)	(0.100)
Pred. male employment		-0.14	-0.0017	-0.11
Lag one		(0.093)	(0.16)	(0.17)
Pred. female employment		0.16*	0.21	0.20**
Lag one		(0.094)	(0.14)	(0.10)
Pred. male employment			-0.15	0.074
Lag two			(0.11)	(0.14)
Pred. female employment			-0.0064	-0.055
Lag two			(0.11)	(0.12)
Pred. male employment				-0.11
Lag three				(0.086)
Pred. female employment				0.062
Lag three				(0.098)
Sum of lags for men	0.15**	0.18***	0.16**	0.24**
	(0.061)	(0.067)	(0.072)	(0.098)
Sum of lags for women	-0.0055	-0.0033	0.051	0.09
	(0.053)	(0.057)	(0.061)	(0.068)
Observations	9,108	8,096	7,084	6,072
Number of municipalities	1,012	1,012	1,012	1,012
State trend	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes

This table shows results of regressions of the log of the fertility rate on the log of predicted formal sector employment (i.e. the reduced form) for men and women and their lags. Standard errors are clustered at the municipality level and reported in parentheses. Regressions are weighted by population of women aged 15-44, and results are for years 2005-2013. Cubic polynomials of the log population of men and women aged 15-44 are included as controls.

Table 1.9: Decomposition of predicted employment into high and low frequency components

	(1)
	Log (GFR)
High-frequency men's employment	0.42**
	(0.18)
High-frequency women's employment	-0.29*
	(0.17)
Low-frequency men's employment	0.14***
	(0.058)
Low-frequency women's employment	-0.0033
	(0.057)
Observations	8,096
Number of municipalities	1,012
State trend	Yes
Pop Controls	Yes
Municipality FE	Yes
Year FE	Yes

This table shows results of regressions of the log of the fertility rate on the decomposition of the log of predicted employment for men and women into high and low frequencies. Standard errors are clustered at the municipality level and reported in parentheses. Regressions are weighted by population of women aged 15-44, and results are for years 2006-2013. Cubic polynomials of the log population of men and women aged 15-44 are included as controls.

Table 1.10: Relationship between gender intensity of sector and wages

	(1) IMSS Male-to-female wage earnings	(2) IMSS Female earnings	(3) IMSS Male earnings	(4) ENOE Male-to-female wage ratio	(5) ENOE Female wage	(6) ENOE Male wage
Proportion male	-0.42***	0.093	-0.27***	-0.17***	-0.054	-0.20***
	(0.042)	(0.090)	(0.088)	(0.052)	(0.075)	(0.066)
Observations	2,206	2,206	2,206	640	640	640
Year FE	Yes	Yes	Yes	Yes	Yes	Yes

Regressions of log wages and log earnings and the wage ratio on proportion of men in each sector, weighted by employment in that sector. Each cell is a sector-year. The ENOE regressions are limited to sectors with a cell size of 10 or greater.

Table 1.11: Impacts on the wage ratio and wages

	(1) Wage ratio	(2) Wage ratio	(3) Wage ratio	(4) Male wage	(5) Male wage	(6) Male wage	(7) Female wage	(8) Female wage	(9) Female wage
Prop. male	-0.19***	-0.19***	-0.060	0.050	-0.16***	-0.036	0.23***	-0.0018	0.034
	(0.055)	(0.051)	(0.054)	(0.069)	(0.056)	(0.069)	(0.079)	(0.060)	(0.075)
Prop. men in high-skill	-0.034		0.44***	0.77***		0.44***	0.87***		0.12
	(0.088)		(0.13)	(0.070)		(0.097)	(0.100)		(0.13)
Prop. women in high-skill		-0.26***	-0.57***		0.70***	0.39***		0.98***	0.89***
		(0.060)	(0.088)		(0.046)	(0.071)		(0.070)	(0.080)
Observations	631	631	628	631	631	628	631	631	628
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Regressions of either wage inequality, log of male wages, or log of female wages in each sector against the proportion of men in each sector and the proportion of each gender in high-skill occupations. Results are for years 2006-2013.

Table 1.12: Impacts on earnings

	(1) Male earnings	(2) Female earnings	(3) Male earnings	(4) Female earnings
Male employment	0.15***	0.17***		
	(0.026)	(0.022)		
Female employment	-0.070***	-0.10***		
	(0.017)	(0.017)		
Pred. male employment			-0.015	0.19***
			(0.068)	(0.061)
Pred. female employment			-0.060	-0.21***
			(0.054)	(0.051)
Observations	9,108	9,108	9,108	9,108
Number of municipalities	1,012	1,012	1,012	1,012
State trend	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Type	OLS	OLS	RF	RF

This table shows results of regressions of log earnings on the log of actual and predicted employment (reduced form) for men and women. Results are for years 2005-2013. Standard errors are clustered at the municipality level.

Table 1.13: First stage regressions for maquiladora employment

	(1) Men	(2) Women	(3) Men	(4) Women
IVmen	0.901***	0.356***	0.855***	0.330***
	(0.0523)	(0.0610)	(0.0536)	(0.0614)
IVwomen	0.212***	0.744***	0.190***	0.709***
	(0.0365)	(0.0596)	(0.0369)	(0.0572)
Observations	38,441	38,441	4,496	4,496
R-squared	0.807	0.764	0.813	0.766
Number of municipalities	2,403	2,403	281	281
Linear state trend	Yes	Yes	Yes	Yes
Sample	Full	Full	Restricted	Restricted
Kleibergen-Paap Wald F statistic	57.06	57.06	51.64	51.64

This table shows results of first stage regressions. Standard errors are clustered at the municipality level and reported in parentheses. Regressions are weighted by population of women aged 15-44, and results are for years 1991-2006. The dependent variables (men and women) represent the total employment of that gender in the municipality divided by the 1990 population aged 15-44 of that gender. The IV represents replaces total employment with total employment solely due to expansions, contractions, openings, and closings of plants. Regressions contain fixed effects for municipality, year, and linear state trends. Cubic polynomials of the log population of men and women aged 15-44 are included as controls. Restricted sample contains only municipalities that have at least 1% employment of men or women in maquiladoras at least one year.

Table 1.14: OLS, IV, and reduced form results using maquiladora employment

	(1) OLS GFR	(2) IV GFR	(3) RF GFR	(4) OLS GFR	(5) IV GFR	(6) RF GFR
Men	0.107	0.259*		0.188	0.334**	
	(0.142)	(0.146)		(0.120)	(0.134)	
Women	-0.0319	-0.154		-0.100	-0.215	
	(0.140)	(0.148)		(0.138)	(0.155)	
IV: men			0.178**			0.215***
			(0.0896)			(0.0691)
IV: women			-0.0597			-0.0889
			(0.0871)			(0.0892)
Observations	38,441	38,441	38,441	4,496	4,496	4,496
Municipalities	2,403	2,403	2,403	281	281	281
Linear state trend	Yes	Yes	Yes	Yes	Yes	Yes
Kleibergen-Paap Wald F statistic		57.06			51.64	
Sample	Full	Full	Full	Restricted	Restricted	Restricted

OLS, IV, and RF regressions. Standard errors are clustered at the municipality level and reported in parentheses. Regressions are weighted by population of women aged 15-44, and results are for years 1991-2006. The independent variables (men and women) represent the total employment of that gender in the municipality divided by the 1990 population aged 15-44 of that gender. Regressions contain fixed effects for municipality, year, and linear state trends. The IV represents replaces total employment with total employment solely due to expansions, contractions, openings, and closings of plants. Cubic polynomials of the log population of men and women aged 15-44 are included as controls. Restricted sample contains only municipalities that have at least 1% employment of men or women in at least one year.

Table 1.15: Impacts on population

	(1) OLS Men	(2) IV Men	(3) OLS Women	(4) IV Women
Male employment (all)	0.023**	-0.36	0.0073	-0.23
	(0.011)	(0.24)	(0.011)	(0.23)
Female employment (all)	0.034***	-0.065	0.046***	-0.14
	(0.013)	(0.16)	(0.013)	(0.16)
Observations	9,108	9,108	9,108	9,108
Number of municipalities	1,012	1,012	1,012	1,012
State trend	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes

OLS and IV regressions for the impact of log male and female employment on log population are shown. Standard errors are clustered at the municipality level.

Table 1.16: Using a Bartik-style IV to identify the impact of employment on fertility

	(1) OLS	(2) IV	(3) RF	(4) OLS	(5) IV	(6) RF
Male emp.	0.0446***	0.48**				
	(0.0143)	(0.20)				
Female emp.	-0.00963	-0.13				
	(0.0129)	(0.11)				
IV (men)			0.20***			
			(0.069)			
IV (women)			-0.0068			
			(0.056)			
Male manuf. emp.				0.0216**	0.22**	
				(0.00885)	(0.099)	
Female manuf. emp.				-0.00552	-0.074	
				(0.00628)	(0.062)	
IV (men) manuf. only						0.11**
						(0.049)
IV (women) manuf. only						-0.036
						(0.050)
Observations	9,108	6,129	6,129	9,108	9,108	9,108
Number of municipalities	1,012	1,012	1,012	1,012	1,012	1,012
State trend	Yes	Yes	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Kleibergen-Paap Wald F statistic		6.24			10.9	

OLS, IV, and reduced form regressions shown. The first three columns include all formal sector employment, while the next set of columns only includes formal sector manufacturing employment. The dependent variable in all cases is the log of the general fertility rate. Standard errors are clustered at the municipality level and reported in parentheses. Regressions are weighted by population of women aged 15-44, and results are for years 2005-2013.

Table 1.17: Impact of male and female employment on fertility

	(1) OLS	(2) IV	(3) RF	(4) OLS	(5) IV	(6) RF
Men	0.0323**	0.30**				
	(0.0140)	(0.13)				
Women	-0.00470	-0.084				
	(0.0129)	(0.078)				
Predicted men			0.14**			
			(0.062)			
Predicted women			-0.0100			
			(0.055)			
Men (ma.)				0.0166**	0.18*	
				(0.00821)	(0.096)	
Women (ma.)				-0.00667	-0.052	
				(0.00594)	(0.051)	
Pred. Men (ma.)						0.094**
						(0.047)
Pred. Wo. (ma.)						-0.026
						(0.043)
Observations	9,108	9,108	9,108	9,108	9,108	9,108
Number of municipalities	1,012	1,012	1,012	1,012	1,012	1,012
State-by-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Pop Controls	Yes	Yes	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Kleibergen-Paap Wald F statistic		9.52			10.3	

OLS, IV, and reduced form regressions for the impact of male and female employment on general fertility rates are shown. Standard errors are clustered at the municipality level and reported in parentheses. Regressions are weighted by population of women aged 15-44, and results are for years 2005-2013. Regressions contain fixed effects for municipality, year, and year by state interactions. Cubic polynomials of the log population of men and women aged 15-44 are included as controls. Cubic polynomials of the log population of men and women aged 15-44 are included as controls.

CHAPTER II

The Impact of Relative Wages on Marriage

2.1 Introduction

Mexico has undergone major social changes in the last three decades. Trade liberalization, starting in the 1980s, led to the decline of import-substituting sectors and accelerated the growth of export-oriented sectors. Recent evidence indicates that these labor market shocks have differentially affected men versus women (Juhn, Ujhelyi, and Villegas-Sanchez, 2014; Garcia-Cuellar, 2001; Artecona and Cunningham, 2002). Indeed, women's labor force participation rates have steadily increased over this period. At the same time, marriage rates have declined, and rates of single motherhood have increased. Can changes in labor markets explain some of the recent trends in family formation in Mexico?[1]

Despite these changes, Mexico continues to lag behind the U.S. and Europe in facilitating women's entry into the labor force, and recent research argues wage discrimination against women, especially low-skilled women in the informal sector, remains common (Popli, 2013). Furthermore, Mexico has a history of conservative gender norms, with the Catholic Church playing a strong role in propagating traditional views of men and women. Notwithstanding these differences in history, institutions, and stages of development, Mexico—and many other

[1] This paper focuses exclusively on changes in marriage between men and women since 1990. A law recognizing same-sex marriage was passed in 2009 and took effect in 2010 in Mexico City and in some other jurisdictions thereafter (but remains prohibited in most localities, though it is required to be recognized everywhere), which unfortunately precludes me from including it in the analysis. Henceforth, the term "marriage" refers only to different-gender couples.

middle-income countries in Latin America—has undergone trends that are similar to those in the U.S. and Europe.

From a policy perspective, these issues matter because Mexico lacks the kinds of government-provided social programs provided by wealthy countries. Instead, family-based social networks effectively function as a social safety net. On the one hand, religion, the lack of a government safety net, and kin-based stigmatization may deter women from avoiding marriage or having a child outside of marriage. On the other hand, social networks and extended households can provide alternative sources of insurance (Fafchamps and Gubert, 2007). The literature on labor markets and marriage has not yet explored how the structural changes in labor markets that have been observed over the last few decades interact with this difference in norms in Mexico and many other developing countries. Other social and economic changes, such as advancements in household technology (Greenwood, Seshadri, and Yorukoglu, 2005), reductions in the costs of obtaining market substitutes for time-intensive commodities that traditionally were produced at home (Sevilla-Sanz, 2005; Ilahi, 2000), declines in the influence of the Catholic Church, and evolving preferences regarding traditional gender norms, have all played an important role in changing Mexicans' views toward marriage.

This paper contributes to a growing literature in labor and family economics that studies the causal effect of changes in labor market opportunities on changes in marriage and family structure (Autor, Dorn, and Hanson, 2017; Bertrand, Kamenica, and Pan, 2015; Shenhav, 2016; Schaller, 2016; Kis-Katos, Pieters, and Sparrow, 2017). In particular, I find that the rise in relative labor demand for women explains at least some of the decline in marriage in Mexico. I also find, unlike recent work focusing on the U.S. (Shenhav, 2016), that these effects are concentrated among higher-educated women. Concentrating on regional variation, I find that urban, wealthier regions of the country have the greatest impacts.

A large literature has focused on how non-wage sources of income affect household bargaining power in developing countries. For instance, Bobonis (2009) studies how an exogenous increase in income for women changes expenditures; Ashraf, Karlan, and Ying (2010) implement a randomized control trial to study how a savings product changes decision-making power within a household; and Duflo (2003) shows that an exogenous increase in women's income, via an old-age pension, changes intra-household spending patterns. Most of this research has focused on bargaining within the household, taking marriage as given, whereas I focus on the household formation decision. Moreover, finding sources of credible, exogenous variation in relative labor demand has proven difficult. Yet, as Anderson and Eswaran (2009) note, earned sources of income may play a much more important role than unearned sources in leading to women's empowerment.

Similar to Aizer (2010), I generate measures of local labor demand for men and women that are based on exogenous demand shocks using a shift-share approach (Bartik, 1991). This is especially important in a setting like Mexico, where fewer than half of all women work. Thus, observed wages are unlikely to serve as a reliable indicator of "potential" wages for all women. Potential wages, not actual wages, determine bargaining power between men and women at their threat point (Pollak, 2005).[2]

The results of this paper provide support for traditional models of the household, where the man specializes in market work and the woman specializes in non-market work (Becker, 1973; Becker, 1981). Comparative advantage, then, leads to gains to marriage through specialization. If women's wages, relative to men's, increase, then the gains to marriage are diminished. To the extent that men and women may have different preferences about whom they marry, when they marry, and the nature of marriage,[3] the results here imply that relative labor market opportunities (and not just unearned income) may also alter bargaining power.

Changes in bargaining power between men and women can have important social di-

[2] The specification of the threat point depends on the model. In early bargaining models, it is specified as divorce, that is, the payoff each individual receives if the marriage ends (Manser and Brown, 1980; McElroy and Horney, 1981). In a less severe formulation, such as in Lundberg and Pollak (1993), the threat point is a noncompetitive equilibrium within marriage where each individual retreats to his or her own "separate sphere." These models focus on bargaining within marriage. Prior to marriage, we can think of a threat point as an outside option, such as staying single.

[3] Pollak (2016) discusses the implications of models that assume whether individuals make binding agreements in the marriage market versus models that allow for bargaining in marriage.

mensions. Latin American countries, including Mexico, have among the highest rates of domestic abuse and rates of violence against women in the world (Waiselfisz, 2015). Increases in labor market opportunities may reduce violence against women by improving their outside option (Aizer, 2010)—that is, allowing them to decline marriage, or delay marriage to search for higher-quality partners. Although I show that increases in relative labor market opportunities for women lead to declines in marriage, I cannot conclude whether these effects are welfare-improving.[4] Declining marriage may lower consumption, increase incidence of poverty, and deprive women of insurance against income shocks. Furthermore, a broad literature has documented negative impacts on children from single motherhood, such as drug use, risky sexual activity, and low test scores, implying that these effects may persist through subsequent generations.[5] Nevertheless, the effects documented in this paper are concentrated mainly among higher-educated women, who command higher wages and have access to greater resources to offset some of these negative effects.

2.2 Conceptual framework

Potential reasons for the decline in marriage include advancements in household technology, changes in social norms and legal institutions, declines in the influence of Catholicism, shortages of "marriageable men" (Edin and Nelson, 2013), and changes in relative labor market opportunities. A broad literature has examined recent trends in marriage and explanations for its decline (e.g. Lundberg and Pollak, 2014 and citations therein).[6] This section presents a simple conceptual framework for describing how a particular indicator of labor

[4] Evidence from Angelucci (2008) suggests that when poor women in rural Mexico obtain higher non-wage income, they receive less abuse from their husbands, which may be welfare-improving for women. Bobonis et al. (2013) find that women who receive transfer benefits are less likely to be victims of physical abuse, though other types of abuse increase. These results also provide an alternative mechanism for how wage or income changes can reduce violence against women, given the potential non-pecuniary costs associated with marriage.

[5] See McLanahan, Tach, and Schneider (2013), McLanahan and Percheski (2008), Autor et al. (2016), and Ayllón and Ferreira-Batista (2015). For evidence that children of less-skilled Hispanic mothers in the U.S. are better off without fathers, see Finlay and Neumark (2010).

[6] Most of this work focuses on the U.S. or Europe. To what extent these trends carry over to a middle-income country like Mexico is a central issue in this paper.

market opportunities—relative wages for women—affects decisions regarding marriage.

I describe two aspects of how relative wages influence a marriage market equilibrium: first, the payoffs to marriage and how these payoffs influence how many individuals choose to marry; and second, selection into marriage, that is, which individuals marry (Pollak, 2016). An individual marries if the utility of marriage exceeds the utility of being single: $U(Marry) > U(Single)$. I assume, as in Becker (1973), that one of the motivations for marriage involves "production complementarities." In Becker's original (1973) formulation, two individuals can generate income together or separately. The output from marriage is the sum of what these individuals produce together. The gains to marriage are larger the more they can produce together, relative to what they would produce on their own. This occurs with economies of scale and when each partner specializes according to his or her comparative advantage. Thus, if women have lower potential wages,[7] men specialize in market work, whereas women specialize in household production.[8] As long as the latter holds, the gains to marriage decrease as women's wages, relative to men's wages, increase. Thus, the utility of marriage is decreasing in the relative wage: $\frac{\partial U(Marry)}{\partial(w_F - w_M)} < 0$.

A second component consists of how relative wage changes affect selection into marriage. That the utility of marriage decreases if the relative wage increases does not imply that all women are equally impacted by a rise in the relative wage. In particular, it is reasonable to believe that the utility of being single is a function of one's own potential wage as well as other factors, such as the level of support one can expect from one's family and social network: $U(Single) = f(w_F, \varepsilon)$. If wages or non-labor sources of income are low enough, and hence $U(Single)$ is low enough, then a decrease in the gains to marriage may not be enough to deter marriage.[9]

[7] Note that what matters here are potential wages, not actual wages. As I show later, women are positively selected into the labor force, and actual wages for the subset of women who work are not substantially lower than wages for men.

[8] An alternative economic way of thinking about the gains to marriage is through consumption complementarities (Stevenson and Wolfers, 2007), rather than through production complementarities.

[9] More formally, this means that $\frac{\partial Prob(Single)}{\partial w_F}$ differs across women, depending on the value of each individual's draw of ε. If for some individuals $U(Married) > U(Single)$ before and after a change in the potential wage, then $\frac{\partial Prob(Single)}{\partial w_F} = 0$.

Despite its simplicity, several predictions emerge from this setup. An increase in the relative wage for women decreases the likelihood of marriage. This occurs because an increase in women's wages reduces the gains to marriage, whereas an increase in men's wages increases the gains to marriage. Moreover, the effect of changes in the relative wage is likely to be blunted among women with the lowest potential wages, as they may not be close to the margin between choosing or foregoing marriage.

2.3 Data and background on wages and marriage

2.3.1 Data description

Data come from the 1990, 2000, and 2010 Mexican censuses.[10] The main analysis focuses on how changes in wages for women, relative to men, affect marriage decisions. Each census uses a different industry classification (the 1990 census uses five-digit sectors, the 2000 census uses three-digit sectors, and the 2010 census uses four-digit sectors). As the identification strategy relies on industrial composition at the local level to identify impacts of relative labor market opportunities on marriage, I need a consistent measure across time to link industries. Hence, I created a concordance table to link industries, aggregating to the 3-digit level.

To exploit variation in local labor markets, I perform the analysis at the municipality-level. Hence, each municipality represents a local labor market in this framework. Since the number of municipalities has grown over time, I use consistent boundary definitions across all years.

2.3.2 Wage gaps

Since this paper exploits changes in relative labor market opportunities due to exogenous demand shocks, it is useful to consider the main sources of the gender wage gap as well as

[10] Data are made available via the Minnesota Population Center. Integrated Public Use Microdata Series, International: Version 6.4 [dataset]. Minneapolis, MN: University of Minnesota, 2015. I do not use earlier years because an earthquake in Mexico City in 1985 destroyed the 1980 census information before it could be digitized.

potential reasons for how and why it has evolved. This subsection briefly reviews a few of the findings from the study of wage gaps.

Research on U.S. labor markets has documented a sizeable wage gap between men and women (Blau and Kahn, 2000; Blau and Kahn, 2006; Blau and Kahn, 2017), with women's relative wages increasing sharply in the 1980s. A large body of work has considered the roles of selection (i.e. cohorts with different characteristics entering the labor force over time, as in Mulligan and Rubenstein, 2008), gender discrimination, gender differences in employment at the occupation-, industry-, and firm-level[11], the impact of children (Waldfogel, 1998), the level of competition in the market (Becker, 1957), and technological shifts (Blau and Kahn, 1997). These explanations are not mutually exclusive. For instance, women may anticipate reducing labor supply in response to having children, which would lead them to lose general or firm-specific capital. At the same time, they may opt for occupations and firms with greater flexibility in working arrangements, even though these may offer lower returns.

Women's wages increased sharply in the 1980s in the U.S., with evidence that positive selection by women into the labor force helped to close the unconditional wage gap (Mulligan and Rubinstein, 2008; Bar, Kim, and Leukhina, 2015). Studies of the gender wage gap over time in Mexico have mainly focused on the impacts of trade liberalization.[12] Formerly a fairly inward-oriented economy, trade reforms began in 1985 in Mexico and culminated in the signing of NAFTA in 1994. This period coincided with moderate increases in labor force participation by women.

Aguayo-Tellez et al. (2010) argue that women's relative wages increased in Mexico during the period of trade liberalization, with both between- and within-industry shifts favoring women. I find similar results. In particular, I evaluate changes in female intensity across

[11] Women are often observed in lower-paying sectors and occupations. Card, Cardoso, and Kline (2016) provide evidence that women are more likely to work in firms with smaller pay premiums.

[12] An important caveat of these studies is that they usually focus on differences-in-differences estimates before and after NAFTA, comparing the border region in Mexico, which is presumably more exposed to trade effects, to other parts of the country. However, many trade reforms preceded NAFTA, and Mexico's economy was undergoing structural changes at the same time, so it is possible the effects of NAFTA are conflated with other policy changes during this period.

industries in my data and find that virtually all sectors had a higher female share in 2010, compared to 1990 (see Figure 1). Garcia-Cuellar (2001), focusing on changes between high- and low-skilled labor, finds a narrowing in the low-skilled gender wage gap that she argues is associated with NAFTA. She argues that these results both confirm the predictions of the Heckscher-Ohlin model because Mexico is abundant in low-skilled female labor, as well as support Becker's (1957) theory because less competitive sectors prior to NAFTA saw greater decreases in the wage gap.

Perhaps surprisingly, given the consistency in international wage gaps in other research (Blau and Kahn, 2003), the unconditional wage gap between men and women in Mexico is small; depending on the way wages are defined, women may even earn more than men.[13] Figure 2 shows average wage gaps (using different ways of constructing the sample) across the three decennial censuses in 1990, 2000, and 2010 for both full-time workers and all workers. Regardless of the manner in which wages are defined, there is a modest increase in relative wages for women in the 1990s, which was followed by a more modest increase or relative stagnation in the 2000s, depending on which measure of wages is used. Of course, given the list of factors associated with the wage gap, these only provide descriptive evidence of the evolution of unconditional wages.

For a closer look at the wage gap, I estimate Mincerian regressions of the natural logarithm of wages for men and women for each of the decennial censuses between 1990 and 2010.[14] Column (1) in Tables 1, 2, and 3 (on data from 1990, 2000, and 2010, respectively) shows the results of regressions that only contain a constant and a male dummy variable, which is not statistically significant from zero in 1990 or 2000 and only marginally positive and statistically significant at the 10% level in 2010. Columns (2) and (3) add extra controls

[13] A similar pattern is documented in Aguayo-Tellez et al. (2010) when including self-employed workers, which I also include in the analysis that follows. Popli (2013) also finds a small differential in raw wages for men and women, whereas conditioning on observable characteristics increases the size of the wage gap. Note that even though the unconditional wage gap is small in the census data, a sizable earnings gap is observed since men work longer hours than women.

[14] These results are only meant to describe wage differences and not to decompose formally the gap into differences due to characteristics of men and women and to differences due to the rewards to these characteristics. See Fortin, Lemieux, and Firpo (2011) for a discussion of these methods.

common in the literature decomposing wage gaps (schoolings, years of experience, years of experience squared, marital status, industry, occupation). As more controls are added, the wage gap grows. Notice, also, that being married is associated with earning higher wages, with married women earning a more significant premium than married men. Since the wage gap grows as more controls are added, women appear to be positively selected into the labor force.[15] Figure 3 supports this view: it shows the mean employment-to-population ratio of women at each year of completed schooling. The relationship between the likelihood of being employed and years of schooling is generally positive. Figures 5 and 6 show mean employment-to-population rates across regions and education groups over time. These graphs indicate consistent increases across all regions and over all skill-types over time, with greater increases in the 1990s, which encompasses many trade reforms, than in the 2000s, which coincides with a long period of stagnation in the Mexican economy.[16]

In sum, the pattern of wage differences between men and women in Mexico observed over the last two decades is as follows: women are positively selected into employment, and employment has increased across both high- and low-skill groups. Overall, the wage gap decreased between 1990 and 2000, but has not narrowed markedly since 2000 due to changes in both supply and demand factors.

2.3.3 Marriage in Mexico

Recent work by sociologists and scholars of Latin America describes women's status and marriage expectations in terms of patriarchy and marianismo[17] giving way to more egalitarian norms in recent times. However, Arrom (1985) argues women's roles did not follow this linear path; instead, she argues that in some ways women had greater autonomy

[15] Note that this implies that that some combinations of controls may be less common among women than among men and vice versa. Ñopo (2008) discusses how to use matching to deal with the lack of a common support in the distribution of attributes among men and women when decomposing gender wage gaps.

[16] Note that these patterns are functions of long-run trends as well as short-run shocks, as these decennial censuses are not taken along the same points in the business cycle.

[17] Marianismo is a term coined by Stevens (1973) as a counterpart to machismo to describe the ideal of femininity in terms of Mary, mother of Jesus. The expectations of marianismo in marriage involve having children and serving one's husband.

in the nineteenth century than in the first half of the twentieth century.[18] She documents high rates of households being headed by women in the nineteenth century, though much of this was due to widowhood. Marriage was traditionally under control of the Catholic Church, but liberal reformers in Mexico and elsewhere in Latin America argued in favor of civil marriage to secularize the institution (Sanders, 2012). Unfortunately, these reforms eliminated equal treatment of the sexes in some conditions. Sanders (2012) provides adultery as an example: Catholic doctrine treated adultery the same for both husbands and wives, but subsequent reforms legalized adultery for husbands yet made it a crime for women. Ultimately, twentieth-century laws have been updated to remove discrimination on the basis of sex, and divorce was enshrined as a legal right in Mexico's constitution of 1917, though conservative local norms and traditions continue to stigmatize it.

Table 4 provides summary statistics showing changes in marriage rates between 1990, 2000, and 2010, and Figure 7 shows the proportion of women who are married by age grouping. Marriage rates among women aged 16-21 have not declined, but marriage rates among women aged 22 and above have steadily declined over these two decades, with the biggest decline among women in their 20s. This is a combination of foregoing and delaying marriage. At the same time, the median age of first marriage has increased over time from twenty years in 1995 to twenty-seven in 2013 (INEGI, 1997; INEGI, 2013). Table 4 also shows that Mexican women are more likely to be heads of households, to live in a household with no children, to be single mothers, and to be divorced in 2010 than in 1990. To the extent that these trends are manifestations of women's increasing bargaining power over this period, the next section focuses on addressing whether these changes are causally related to changes in relative labor demand for women.

[18] Arrom (1985) also notes that Mexico granted women greater property rights during this time than the U.S.

2.4 Identification of relative labor demand shocks for women

2.4.1 Relative bargaining power in the labor market

The principal problem with identifying the impact of relative bargaining power in the labor market is obtaining a credible proxy for an individual's bargaining power. I build on two important insights in Pollak (2005). First, I use a proxy for labor demand to isolate exogenous demand shocks for wages for men and women. I do not use not earnings or hours worked to proxy for bargaining power. Higher earnings may simply signify working more hours (and hence devoting less time to household production and leisure), but that does not imply having more bargaining power relative to one's spouse. Having a higher wage, however, does indicate having more bargaining power in the labor market.

The second major point in Pollak (2005) is determining how to specify an individual's wage. In typical bargaining models, bargaining is determined by well-being at one's "threat point", not at the observed equilibrium. For instance, a woman may choose not to work if she marries, in which case her wages are not observed in equilibrium. Suppose that if she were to remain single, however, she would work; this is her potential or outside wage option.

Own-wages cannot serve as a reliable proxy for relative bargaining power. Prevailing local wages might serve as a better indicator of relative bargaining power in the labor market, but these still suffer from being observed wages in equilibrium. In particular, when female labor force participation is low and unrepresentative of working-age women, shocks to local wages may reflect both supply and demand decisions of women. For instance, if a local area experiences increases in labor force participation among less-skilled women, prevailing wages may decrease, even though potential wages may have increased for these women.

To identify exogenous sources of variation in an individual's potential wage, I create an index for local labor demand following the methodology in Bartik (1991), Blanchard and Katz (1992), and many subsequent papers. Katz and Murphy (1992) formalize this measure and relate it to wages in the context of a supply and demand framework.[19] The index is defined as follows:

$$D_{m,g,y} = \sum_{ind} (\theta_{m,g,ind})(Emp_{-m,g,ind,y})$$

$$\text{where } \theta_{m,g,ind} = \frac{Emp_{m,g,ind,y=1990}}{Emp_{m,g,y=1990}}.$$

The measure of local labor demand is calculated for each gender g in each municipality m in each year y by weighting industry-wide (log) employment for each gender g in each year by the share of men or women employed in a given industry ind in that municipality and then summing over all industries.[20] Specifically, the numerator in $\theta_{m,g,ind}$ is equal to the employment of group g in municipality m in industry ind in 1990, and the denominator in the fraction is equal to the total employment of group g in municipality m in 1990. This employment share does not vary across years to prevent endogenous responses in labor supply from biasing the measure. Here, $Emp_{-m,g,ind,y}$ is the (log) employment in industry ind. I also follow the common practice of excluding municipality m from the calculation of $Emp_{-m,g,ind,y}$ to prevent local employment changes from influencing the measure; this is indicated by $-m$. I take the difference between women's labor market demand and men's labor market demand as my measure of relative bargaining power in the labor market.

2.4.2 Labor demand variation

Before moving on to the econometric specification and the results, it is useful to consider what explains the variation in labor demand for women versus men. Similar to work studying demand for skilled versus unskilled workers, we can think of men and women as being imperfect substitutes in firms' production functions. Over the time period studied here, the proportion of the labor force that was female rose, which potentially reflects increasing labor supply among women, increases in labor demand, or both. It is also important to

[19] A similar discussion, as well as references, can be found in Katz and Autor (1999).

[20] Bartik (1991) discusses the conceptual rationale for using weighted industry shares as a measure of the local demand for labor.

note that Mexico experienced within-industry shifts, where women joined male-dominated industries. Indeed, nationwide growth in almost all industries was higher for women than for men, implying that that sectors became feminized over time. However, the correlation between female and male growth within industries was very high over this period, indicating that within-industry shifts were not the main driver of changes in relative labor demand for women.

The measure of labor demand for each gender exploits the fact that different industries employ men and women in different proportions. We would observe an increase in relative labor demand for women if between-industry shifts in labor demand, such as from sector-specific productivity shocks or from changes in product demand (either arising from domestic sources or from international trade), led industries that predominantly employed men to decline (or grow relatively less) than industries that employed relatively more women. Indeed, I find that industries that employed relatively more women initially experienced larger growth. These findings mirror those in Aguayo-Tellez et al. (2010), who find a re-allocation in employment across sectors in favor of women in Mexico. They argue that policies connected to trade liberalization led to a shift in labor demand toward industries that disproportionately employed women and raised their wages.

Evaluating the sector-level census data used in this paper, I find that for sectors that disproportionately employed men, the main sources of decline in employment shares were concentrated in a subset of sectors in agriculture, forestry, fishing, hunting, and manufacturing. The latter includes industries in the manufacturing of metal industries, machinery and equipment, and non-metallic minerals. These sectors also suffered among the largest wage declines in the data. Previous work has documented the impact of trade liberalization, including the passage of NAFTA in 1994, as triggering the decline in these sectors (Nicita, 2004; Hanson and Harrison, 1999; Revenga, 1997).

For women, sources of employment growth in manufacturing included the clothing industry and the food industry (especially in the 1990s), but in general growth in sectors that employed disproportionately more women was more varied and mainly concentrated outside of manufacturing. Women's employment increased in a variety of sectors in wholesale and retail commerce. Growth in household employees in private households, one of the biggest employers of women, facilitated the entry of less-skilled women into the workforce. In addition to being one of the biggest sources of employment growth, this industry also experienced among the largest wage growth between 1990 and 2010 (a difference in average log wages of 0.22). This indicates that employment growth in female-centric sectors was not merely due to a secular increase in labor supply among women. Other sectors that were disproportionately likely to employ women include services in real estate, finance, and other professional services. This growth facilitated the entry of more-skilled women into the workforce in major cities. These trends parallel those in the U.S. and many other countries.[21]

These labor demand shifts were not evenly dispersed across the country. Growth in women's employment in small-scale manufacturing (clothes, food) occurred predominantly in the border region,[22] as well as the central belt and Mexico City regions. Growth in services was concentrated in the major cities of the same areas. To summarize, the growth in a subset of manufacturing sectors and in services drove the increase in the women's employment in the border, capital, and center regions, with large increases in women's labor force participation since 1990. Men's employment grew relatively more slowly, with negative shocks to male-dominated sectors in agriculture throughout rural parts of the country, and negative shocks to manufacturing in the border, capital, and center regions.

[21] Rendall (2017) and Pitt, Rosenzweig, and Hassan (2012) provide a theoretical foundation for a shift in labor market demand in favor of "brain" versus "brawn", where women are assumed to have a comparative advantage in "brain".

[22] Much of this growth came from the expansion in maquiladoras, which increased labor market opportunities for women and were mainly located in the border.

2.5 Estimating the impact of relative bargaining power on marriage

To estimate the impact of changes in relative labor demand for women on marriage and related outcomes, I estimate variants of the following reduced-form specification for Mexican municipalities:

$$y_{m,y,a} = \lambda + \beta RBP_{m,y} + \boldsymbol{\alpha F(X)}_{m,y,a} + \gamma_m + \delta_y + \theta_a + \varepsilon_{m,y,a}. \tag{2.1}$$

I collapse observations into municipality-age-year-education cells, where each observation represents a woman of age a in a given census year y living in municipality m. RBP is relative bargaining power in the labor market, as described earlier. I include municipality fixed effects, which control for unobservable differences common to municipalities, age fixed effects, which control for differences in marriage propensity across ages, and year fixed effects, which control for nationwide shocks to marriage. In addition, some specifications include education fixed effects, year by education interactions, age by education interactions, and quadratic polynomials in the proportion of the local population that is male in $\boldsymbol{F(X)}$. I use four education categories: individuals with less than primary schooling completed, individuals who completed primary school, individuals who completed lower secondary schooling, and individuals who completed secondary schooling or more.

The demographic controls are designed to control for the effects of the sex ratio, which may have an important, independent role in influencing marriage decisions (Abramitzky, Delavande, and Vasconcelos, 2011; Angrist, 2002; Chiappori, Fortin, and Lacroix, 2002). To capture the effects of the relevant demographics, I compute local populations of men and women at each age and use the male proportion of the population for a given woman's age as the relevant control, which provides a more precisely estimated effect of the relevant sex ratio than a coarser measure across a wider age interval.[23] All regressions are limited to

[23] Staggering the relevant sex ratio, i.e. identifying the number of men for slightly younger women, leads to very similar results, so I use same-age proportions for simplicity.

cells with non-missing observations for men and women in all ages between 22-44. I use census-provided weights. Standard errors are clustered at the municipality-level.

2.5.1 Results of the estimation

The results from the main specification can be found in Panel A in Table 5. Column (1) includes only municipality, year, and age fixed effects. In columns (2) and (3) I add more of the controls described in equation (1). The estimate of β changes little, with the specification in column (3) indicating that a 10% increase in relative labor market demand leads to a 1.3% decline in the probability of being married. Using the actual change in relative labor demand and the mean change in marriage rates, I find that these estimates imply that about a quarter of the decline in marriage can be explained by the change in relative labor market opportunities for women.

To disentangle whether changes in labor demand for men or women are driving the results, I separate the shift-share index into its representative parts by gender. Results are presented in Panel B of Table 5. Column (3), which includes the most controls, indicates that a 10% increase in potential wages for women, conditional on men's potential wages, decreases the probability of a woman being married by 0.9%. A 10% increase in potential wages for men, conditional on women's potential wages, increases the probability of marriage by 1.5%. These results are consistent with the theory: increases in men's labor demand, and hence increases in their potential wages, amplify the attractiveness of marriage, whereas improvements in women's labor demand decrease the gains to marriage. In the specification with the most controls, an F Test rejects the null hypothesis of equality of the magnitude of the coefficients at the 5% level.

The magnitude of the male coefficent is larger. One possible reason for this is that the effects of the labor demand proxy are driven by men's wages, not women's wages. This makes sense especially in a setting where fewer than half of all women work, such that women's wages may play a less important role in household decisions. Morever, even if a woman does not expect to work, a decline in men's wages may lead her to delay, if not abstain from, marriage. Indeed, the results in Table 7 show the biggest impacts at the youngest ages and a weaker effect at higher ages. Another possibility for the difference in estimates of men's and women's potential wages is individual heterogeneity. It is possible that the labor demand proxy does a better job of predicting men's wages since men have uniformly high labor force participation rates, whereas women who work comprise a self-selected sample that may not be as representative of women's outside option.

2.5.2 Other household outcomes and heterogeneity in effects

In this section I evaluate the impact of changes in relative labor demand on other relevant household formation outcomes. Results are shown in Table 6. Column (1) shows the results of estimating the impact of the relative labor demand index on divorce. The coefficient is close to zero, indicating that increases in labor demand for women's do not appear to increase exit from marriage. Column (2) indicates that a 10% increase in the relative labor demand index increases the probability that a woman is a head of the household by 0.37%. Turning to fertility-related outcomes, column (3) shows that a 10% increase in relative labor demand for women increases the probability of there being no children in the household by 1.1%. To focus on an outcome commonly studied in the U.S and other developing countries but less well-studied in the developing countries, I look at single motherhood and find a small increase in the probability of single motherhood.[24] Finally, I look at the impact on the probability of being a widow and find no statistically significant effect. Macabre theories aside, there is no reason to think that relative wages would affect widowhood status, so this may be seen as a falsification test.[25] Together, these results indicate that an increase in

[24] Single motherhood increased from 10% in 1990 to 16% in 2010. Mexico has provided family planning since 1974, when a law establishing free family planning services was passed, though access to clinics providing these services, especially in rural areas, remains an issue. Abortion is regulated at the state level and remains very rare in Mexico. Since 2007, abortion has been available to women living in Mexico City (in limited circumstances), and some, but not all, states have followed in liberalizing their abortion laws.

[25] There is a mechanical correlation between widowhood and marriage in the sense that one must have been married to be widowed, so relative wage options can influence widowhood via their effect on marriage. Nevertheless, the number of widows in this age range is very small, so such an effect would be difficult to

women's relative labor demand does not only lead reduce the likelihood of marriage; it also causes a fundamental change in status within the household, including changes in fertility. These results are consistent with other work that shows that non-labor sources of income, when directed toward women and away from men, increase their bargaining power in the household (e.g. Duflo, 2003; Lundberg, Pollak, and Wales, 1997).

Table 7 shows results by age grouping. The results are not surprising: the largest impacts are among women in their twenties, i.e. those most likely on the margin of considering marriage. Impacts become progressively smaller among women past age 29, and the impact is close to zero among women ages 50-59.

Turning to regional impacts, Table 8 shows the results of estimating the main specification separately for municipalities in the border, north, center, capital, and south of the country. The largest results are in the border region with the U.S., and the weakest results are in the south, which is more rural and agricultural than other parts of the country. Separating individuals by urban and rural status shows much larger results in magnitude among urban residents (results not shown here). These results suggest that the biggest impacts are among wealthier, more educated, and more urbanized regions. These results are consistent with the sources of growth in relative labor demand for women: regions that had a higher concentration of female-centric sectors that subsequently saw larger increase in employment also saw greater decreases in the likelihood of marriage.

Table 9 shows the results of estimating equation (1) separately for four educational categories. The magnitude of the estimated impact of changes in relative labor demand for women generally increases with a woman's level of education. These results are consistent with the earlier results in Table 5 and with the conceptual framework suggested earlier. If women's potential wages are sufficiently low, they may not be on the margin between choosing or foregoing marriage because an increase in relative wages is not enough for them to earn an independent living. In this framework, the relative wage is most relevant for women

detect.

who can earn a high enough wage to be able to delay or abstain from marriage.

It is important to note that although the results on heterogeneity across educational groups are consistent with theory, the theoretical justification is not the only reason why results are weaker for women with less education. An alternative explanation follows from the identification strategy employed here: an ideal measure of relative labor market opportunities would differ across all observable and unobservable characteristics to capture each individual's outside option. The shift-share proxy for labor demand, by construction, has the virtue of exploiting all demand shocks that differentially affect employment across sectors, as opposed to focusing on demand shocks solely due to a single mechanism, such as particular policies (e.g. a reduction in tariffs). Yet there is no reason to believe that labor market shocks across industries affect all individuals equally. First, the types of shocks across industries may differentially affect individuals of varying types, insofar as industries differ in which skill types they hire. [26] Second, if less-educated women have more elastic labor supply, then increases in demand may not translate to large increases in wages. Third, research on Mexican labor markets and the gender wage gap has found that the unexplained wage gap is larger for less-skilled women than for more-skilled women, which may be evidence of greater wage discrimination against less-skilled women (Popli, 2013). All of these may lead to a finding of heterogeneous effects even in the absence of true heterogeneity.

2.6 Conclusion

As in the U.S. and much of Europe, women in Latin America today are more likely to be single, to delay getting married, to head a household, and to be single mothers. However, outside of the U.S., the extent to which changes in women's employment and wages have affected these trends has not been studied. This paper exploits exogenous shocks to labor demand across Mexican municipalities to identify sources of bargaining power in the labor

[26] For instance, suppose demand for the output of the financial services sector increases. This increase in demand may not translate into a meaningful increase in potential wages for an unskilled woman living in a municipality with a high employment base in financial services

market. Using the resulting measure of women's labor demand relative to men's labor demand, I show a causal link between women's relative labor market opportunities and family formation outcomes. These results provide support for the theories of new home economics, which argue that that a rise in women's potential wages reduces the traditional gains to marriage due to specialization.

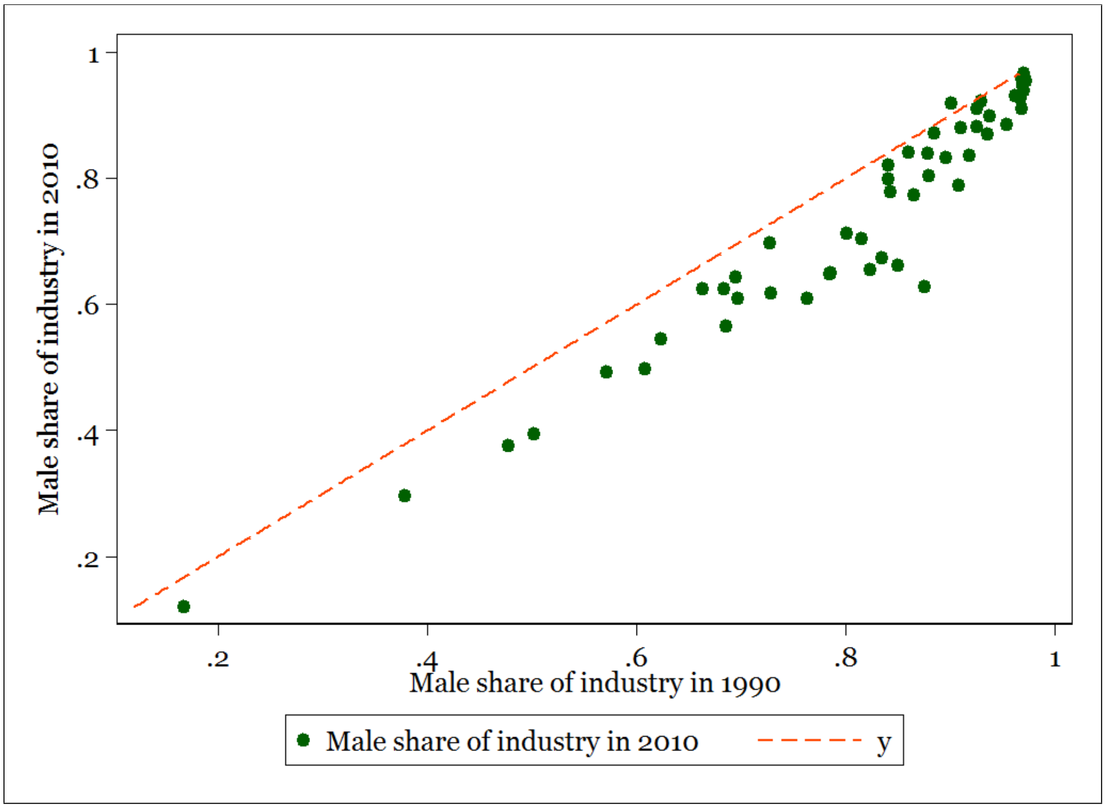
It is also possible that the effects observed here represent "disequilibrium effects."[27] In that case, it remains an open question as to whether the declines in marriage observed thus far continue, or whether an evolution in gender norms allows a turnaround in family formation.

While Mexico, along with much of Latin America, has followed a similar demographic trajectory as the U.S. and Western Europe, it nevertheless has failed to develop a welfare system providing unemployment insurance and social support. The findings of this paper indicate that the rise in relative labor demand has increased women's bargaining power and has led to a decline in marriage. At the same time, increasing numbers of single individuals without the benefit of a social safety net may also lead to more individuals being vulnerable to poverty. The connection between labor markets and changing family patterns, then, should be of broad interest to policymakers.

[27] Pollak (2016) notes that women's greater role in market work has disrupted traditional expectations about gender norms in the home. In the long run—that is, longer than the two ten-year differences observed here—norms may catch up to labor market realities, and the observed relationship between labor market opportunities and marriage may dissipate. For evidence on this point with respect to fertility rather than marriage, see García-Manglano, Nollenberger, and Sevilla (2014). For a broader discussion on how changes in gender relationships affect fertility and marriage, see Goldscheider, Bernhardt, and Lappegård (2015).

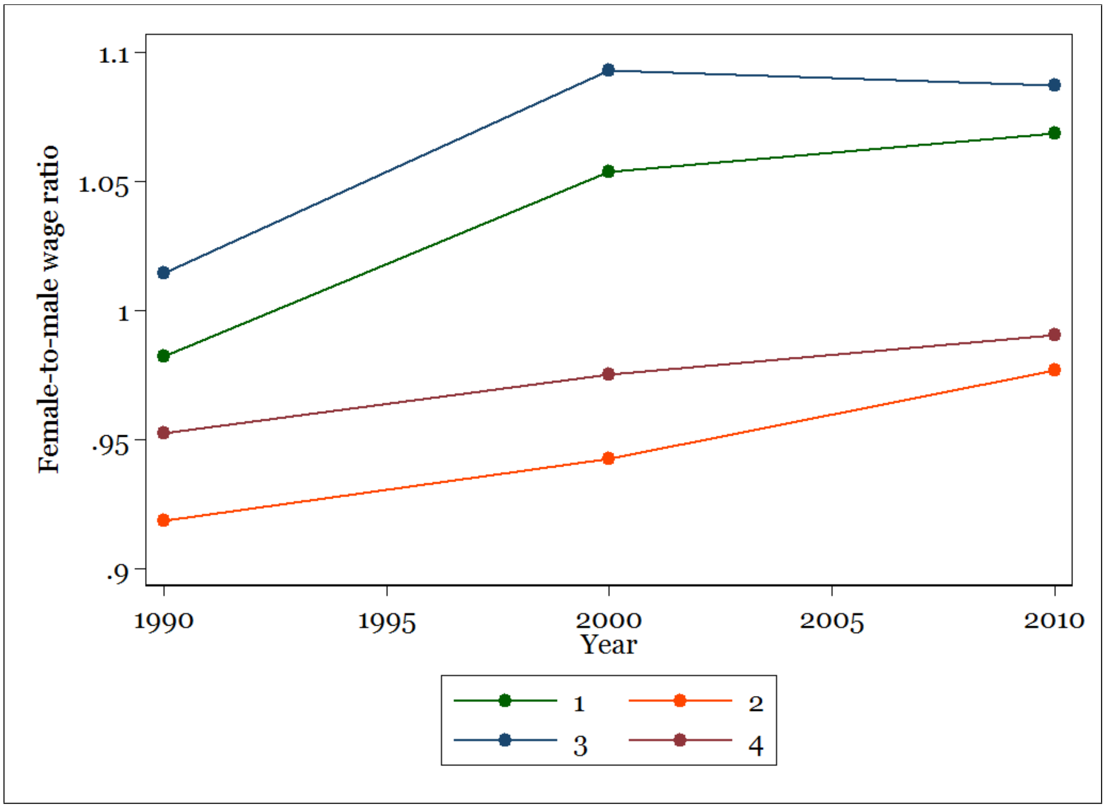
2.7 Figures

Figure 2.1: Male share of 3-digit sectors between 1990 and 2010



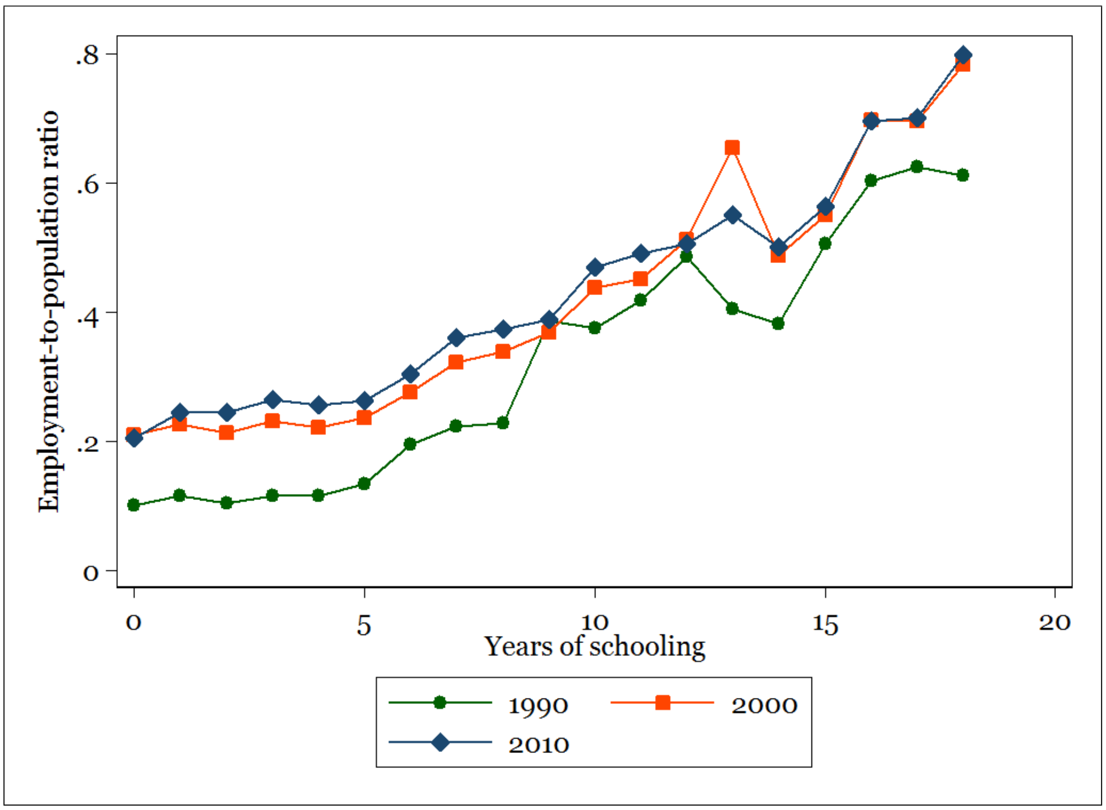
Notes: Industry groupings are at the 3-digit level. The 45 degree line is shown for comparison. Data come from the Mexican Censuses 1990-2010.

Figure 2.2: Female to male wage ratio over time



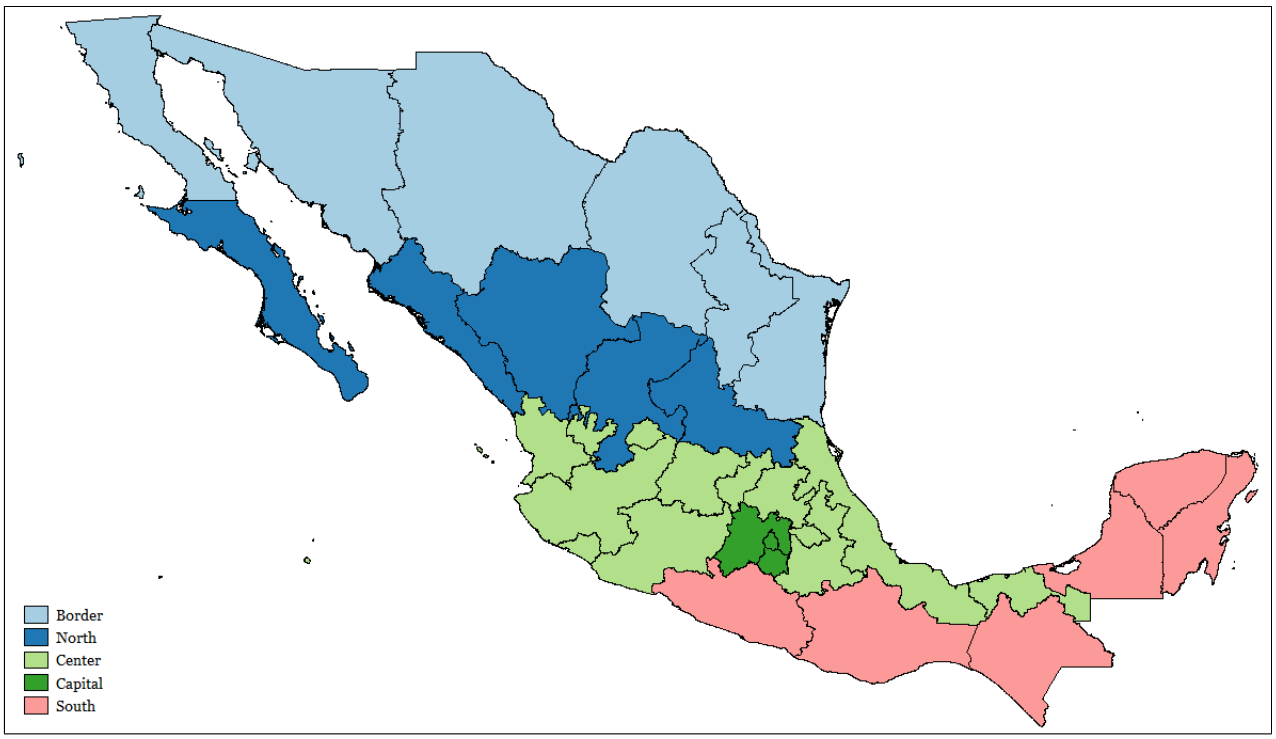
Notes: Wage ratio 1 shows the hourly wage for all with positive hours, after trimming both ends at 1%. Wage ratio 2 shows the hourly wage for all with more than 30 hours, after trimming both ends at 1%. Wage ratio 3 shows the hourly wage for all with positive hours, after trimming both ends at 2%. Wage ratio 4 shows the hourly wage for all with more than 30 hours, after trimming both ends at 2%. Data come from the Mexican Censuses 1990-2010.

Figure 2.3: Mean employment-to-population ratio of women across years of schooling



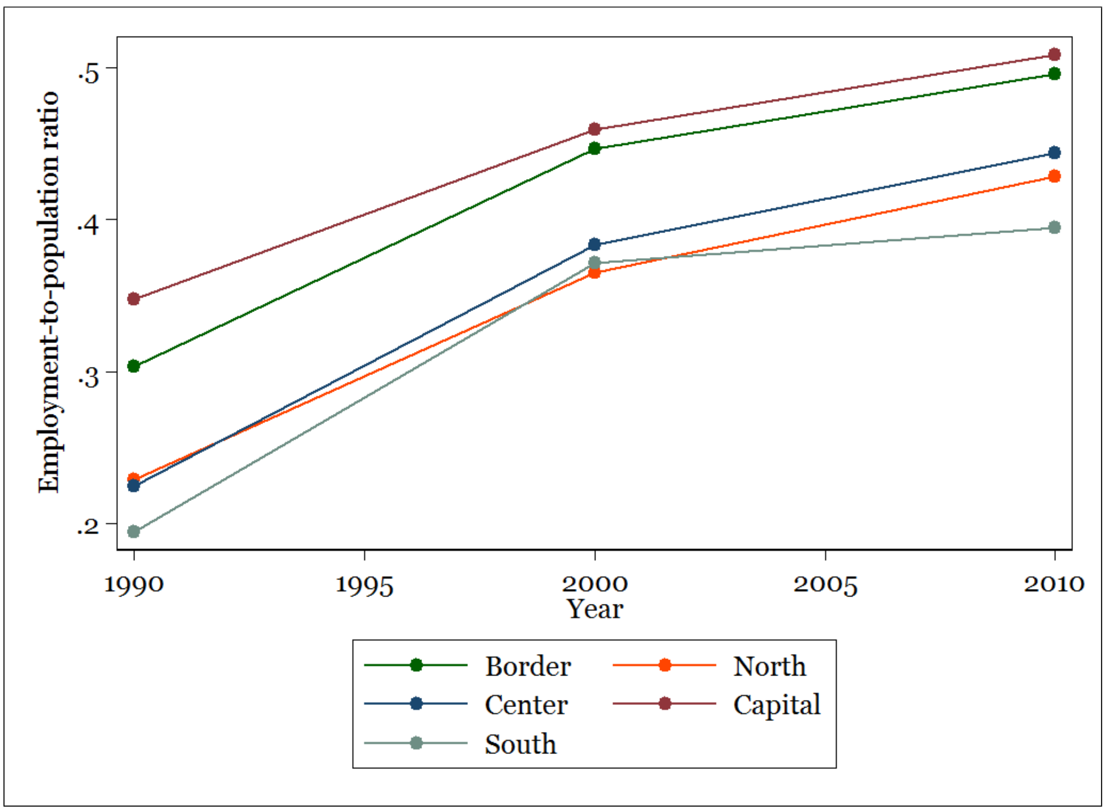
Notes: Each point represents the mean employment-to-population ratio of women aged 22-44 for women with a given number of years of schooling. Only individuals reporting positive income and hours are counted as employed. Individuals reporting more than 18 year of schooling are all grouped into the category "18." Data come from the Mexican Censuses 1990-2010.

Figure 2.4: Map of Mexico



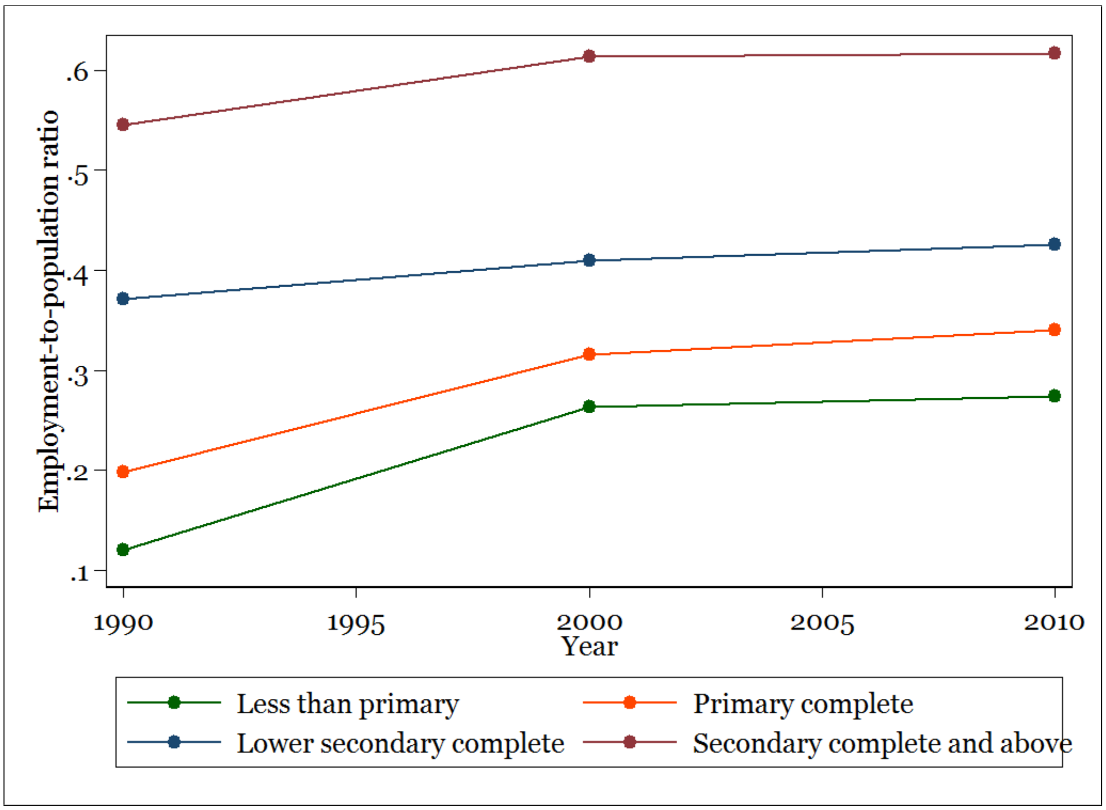
Notes: This graph displays a map of Mexican federal entities according to region.

Figure 2.5: Mean employment-to-population rates of women across regions in Mexico



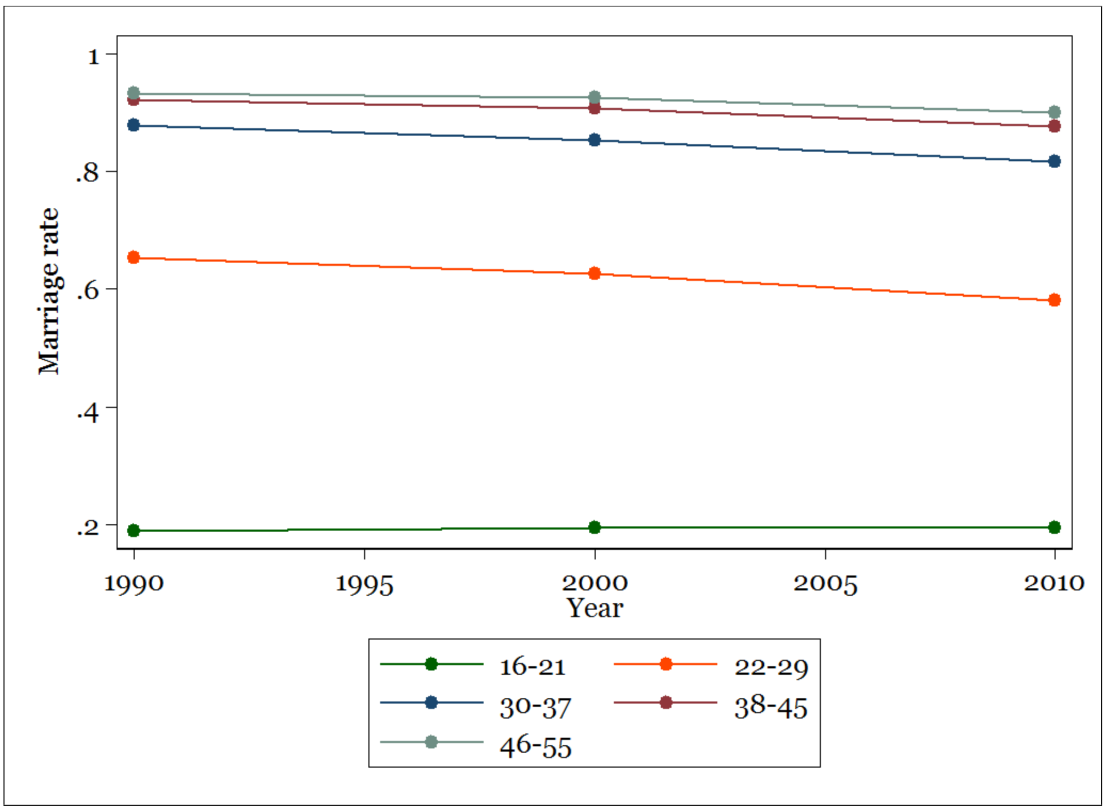
Notes: See Figure 4 for which states are included in which region. The graph shows mean employment-to-population rates in each region across time for women aged 22-44. Data come from the Mexican Censuses 1990-2010.

Figure 2.6: Mean employment-to-population rates of women across educational groups in Mexico



Notes: The graph shows mean employment-to-population rates for groups with different levels of schooling. The lowest group only includes women with less than primary schooling; the next group only includes women with primary schooling; the next group includes women with lower secondary completed; and the last group includes women with secondary schooling complete or more. Data come from the Mexican Censuses 1990-2010.

Figure 2.7: Mean marriage rates by age group



Notes: The graph shows mean marriage rates for women in different age groups. Data come from the Mexican Censuses 1990-2010.

2.8 Tables

Table 2.1: Individual wage regression results

	Log (wage)	Log (wage)	Log (wage)	Log (wage)
	(1)	(2)	(3)	(4)
Male	0.015	0.048**	0.058**	0.12***
	(0.026)	(0.020)	(0.024)	(0.021)
Experience		0.040***	0.040***	0.025***
		(0.0014)	(0.0014)	(0.00071)
Experienceˆ2		-0.00060***	-0.00060***	-0.00035***
		(0.000026)	(0.000025)	(9.0e-06)
Years of schooling		0.077***	0.077***	0.047***
		(0.0031)	(0.0028)	(0.0012)
Male*years of schooling			-0.0011**	-0.0013***
			(0.00050)	(0.00021)
Married				0.10***
				(0.013)
Married*male				-0.040***
				(0.014)
Observations	1,620,918	1,620,918	1,620,918	1,544,595
R-squared	0.000	0.190	0.190	0.288
Industry and Occupation dummies	No	No	No	Yes

Standard errors are clustered at the state level and reported in parentheses. Coefficients are from OLS regressions, where the dependent variable is the natural logarithm of the wage in 2010 pesos. Data come from the 1990 Population Census. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.2: Individual wage regression results

	Log (wage) (1)	Log (wage) (2)	Log (wage) (3)	Log (wage) (4)
Male	0.014	0.10***	0.27***	0.23***
	(0.025)	(0.017)	(0.026)	(0.021)
Experience		0.032***	0.033***	0.023***
		(0.00080)	(0.00080)	(0.00056)
Experienceˆ2		-0.00039***	-0.00040***	-0.00030***
		(0.000019)	(0.000019)	(0.000010)
Years of schooling		0.11***	0.12***	0.078***
		(0.0036)	(0.0031)	(0.0028)
Male*years of schooling			-0.019***	-0.0098***
			(0.0016)	(0.0014)
Married				0.13***
				(0.011)
Married*male				-0.046***
				(0.010)
Observations	2,064,853	2,051,924	2,051,924	1,921,177
R-squared	0.000	0.348	0.351	0.430
Industry and Occupation dummies	No	No	No	Yes

Standard errors are clustered at the state level and reported in parentheses. Coefficients are from OLS regressions, where the dependent variable is the natural logarithm of the wage in 2010 pesos. Data come from the 2000 Population Census. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.3: Individual wage regression results

	Log (wage)	Log (wage)	Log (wage)	Log (wage)
	(1)	(2)	(3)	(4)
Male	0.035*	0.13***	0.34***	0.28***
	(0.021)	(0.014)	(0.025)	(0.015)
Experience		0.025***	0.026***	0.019***
		(0.00077)	(0.00076)	(0.00049)
Experienceˆ2		-0.00030***	-0.00031***	-0.00025***
		(0.000014)	(0.000014)	(8.2e-06)
Years of schooling		0.087***	0.10***	0.064***
		(0.0028)	(0.0022)	(0.0020)
Male*years of schooling				-0.014***
				(0.00099)
Married				0.10***
				(0.0080)
Married*male				-0.043***
				(0.0095)
Observations	2,238,244	2,227,491	2,227,491	2,048,993
R-squared	0.001	0.254	0.259	0.356
Industry and Occupation dummies	No	No	No	Yes

Standard errors are clustered at the state level and reported in parentheses. Coefficients are from OLS regressions, where the dependent variable is the natural logarithm of the wage in 2010 pesos. Data come from the 2010 Population Census. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.4: Summary statistics

	1990		2000		2010	
	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
Married proportion	0.755	0.430	0.720	0.449	0.684	0.465
Divorced proportion	0.037	0.189	0.062	0.241	0.083	0.276
Female head of household	0.088	0.284	0.106	0.307	0.123	0.328
Proportion with no children in household	0.186	0.389	0.205	0.404	0.215	0.411
Proportion who are single moms	0.088	0.283	0.115	0.319	0.147	0.354
Urban proportion	0.796	0.403	0.819	0.385	0.817	0.387
Log relative female/male wage	-0.073	0.186	-0.057	0.132	-0.064	0.091
Educational composition						
Less than primary	0.335	0.472	0.188	0.391	0.102	0.303
Primary completed	0.269	0.443	0.256	0.437	0.193	0.394
Lower secondary completed	0.197	0.398	0.264	0.441	0.319	0.466
Secondary complete and above	0.200	0.400	0.292	0.455	0.386	0.487

Summary statistics for women aged 22-44 in the Mexican Population Census are shown. Census-provided weights are used.

Table 2.5: Impact of relative labor demand on marriage

	(1) Married	(2) Married	(3) Married
Panel A			
Relative labor demand	-0.11***	-0.12***	-0.13***
	(0.020)	(0.021)	(0.025)
Panel B			
Female labor demand proxy	-0.086***	-0.080***	-0.090***
	(0.022)	(0.022)	(0.023)
Male labor demand proxy	0.12***	0.13***	0.15***
	(0.025)	(0.025)	(0.027)
F test (p value)	0.1809	0.0640	0.0348
Observations	4,099,751	4,099,751	4,099,751
Municipality FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Age FE	Yes	Yes	Yes
Education Controls	No	Yes	Yes
Demographic Controls	No	No	Yes

Standard errors are clustered at the municipality level and reported in parentheses. The dependent variable is a dummy for whether the respondent is married. Panels A shows the impact of relative labor demand, defined as the difference between the index for female labor demand and male labor demand, as described in the text. Panel B shows the impact of shift-share index for each gender separately. The p-values of an F test of the equality of the magnitude of the coefficients on the male and female wage proxies in Panel B are also shown. Each unit of observation is a woman aged 22-44. Census-provided weights are used. Education controls include four categories for schooling, and are entered individually and interacted with the year in the second column. The third column adds education x age fixed effects and population controls, which include a quadratic polynomial in the local proportion of men in each individual's age. Data come from the 1990, 2000, and 2010 Population Censuses. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.6: Impact of relative labor demand on family-related outcomes

	(1) Divorced	(2) Female head of household	(3) No children in household	(4) Single mom	(5) Widowed
Relative labor demand	0.0079	0.037***	0.11***	0.026*	-0.0029
	(0.0096)	(0.011)	(0.018)	(0.013)	(0.0026)
Observations	4,099,751	4,042,810	4,018,083	4,018,083	4,099,751
Municipality FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Age FE	Yes	Yes	Yes	Yes	Yes
Education Controls	Yes	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes	Yes

Standard errors are clustered at the municipality level and reported in parentheses. The dependent variable (shown in the column heading) is a dummy variable. Each unit of observation is a woman aged 22-44. Census-provided weights are used. Education controls include four categories for schooling, and are entered individually and interacted with the year in the second column. The third column adds education x age fixed effects and population controls, which include a quadratic polynomial in the local proportion of men in each individual's age. Data come from the 1990, 2000, and 2010 Population Censuses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 2.7: Impact of relative labor demand on marriage by age grouping

	(1) Married Ages 16-21	(2) Married Ages 22-29	(3) Married Ages 30-39	(4) Married Ages 40-49	(5) Married Ages 50-59
Relative labor demand	-0.13***	-0.20***	-0.083***	-0.067***	0.029
	(0.025)	(0.029)	(0.022)	(0.023)	(0.022)
Observations	1,543,961	1,722,568	1,722,545	1,191,228	793,024
Municipality FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Age FE	Yes	Yes	Yes	Yes	Yes
Education Controls	Yes	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes	Yes

Standard errors are clustered at the municipality level and reported in parentheses. The dependent variable is a dummy for whether the respondent is married. Each unit of observation is a woman in a given age interval. Census-provided weights are used. Education controls include four categories for schooling. Controls include education, education x year, education x age fixed effects and population controls, which include a quadratic polynomial in the local proportion of men in each individual's age. Data come from the 1990, 2000, and 2010 Population Censuses. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.8: Impact of relative labor demand on marriage across regions

	(1) Border	(2) North	(3) Center	(4) Capital	(5) South
Relative labor demand	-0.16***	-0.096**	-0.11***	-0.11**	-0.032
	(0.055)	(0.046)	(0.027)	(0.048)	(0.025)
Observations	658,226	345,687	1,630,434	944,933	520,471
Municipality FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Age FE	Yes	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes	Yes
Education Controls	Yes	Yes	Yes	Yes	Yes

Standard errors are clustered at the municipality level and reported in parentheses. The dependent variable is a dummy for whether the respondent is married. Each unit of observation is a woman aged 22-44. Census-provided weights are used. Each column shows results for women living in the region in the column heading. Education controls include four categories for schooling. Controls include education, education x year, education x age fixed effects and population controls, which include a quadratic polynomial in the local proportion of men in each individual's age. Data come from the 1990, 2000, and 2010 Population Censuses. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.9: Impact of relative labor demand on marriage by educational grouping

	(1) Married	(2) Married	(3) Married	(4) Married
Relative labor demand	-0.052***	-0.11***	-0.18***	-0.21***
	(0.019)	(0.025)	(0.032)	(0.035)
Observations	931,380	1,056,583	1,060,187	1,051,601
Municipality FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Age FE	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes

Standard errors are clustered at the municipality level and reported in parentheses. The dependent variable is a dummy for whether the respondent is married. Each unit of observation is a woman aged 22-44. Census-provided weights are used. Each column divides women into a separate education category: column 1 only includes women with less than primary schooling; column 2 only includes women with primary schooling; column 3 includes women with lower secondary completed; and column 4 includes women with secondary schooling complete or more. Controls include education, education x year, education x age fixed effects and population controls, which include a quadratic polynomial in the local proportion of men in each individual's age. Data come from the 1990, 2000, and 2010 Population Censuses. *** p<0.01, ** p<0.05, * p<0.1.

CHAPTER III

The Wage Distribution, Local Labor Market Outcomes, and the Chinese Import Shock

3.1 Introduction

China's transition from central planning to a market economy and its integration into world markets has had a striking impact on international trade. During the same period, manufacturing employment has fallen and wage inequality has risen in many countries. The simultaneity of these events has led to a contentious debate among policymakers and the public regarding to what extent these trends are linked. It has also sparked a large and growing literature among economists, with recent work on how China has shaped local economies[1], the time-paths of wages[2], job polarization[3], marriage and divorce[4], prices[5], innovation[6], and plant entry and exit.[7]

More broadly, a large body of work has studied the impact of trade shocks, including tariff cuts, currency crises, offshoring, and trade penetration from developed countries, on

[1] Autor et al. (2013) study local labor markets in the U.S., and Donoso, Martin, and Minondo (2015), Balsik, Jensen, and Salvanes (2015), Malgouyres (2017), and Dauth, Findeisen, and Suedekum (2014) apply similar methods in Spain, Norway, France, and Germany, respectively.

[2] See Autor, Dorn, Hanson, and Song (2014).

[3] See Keller and Utar (2016).

[4] See Autor, Dorn, and Hanson (2017).

[5] See Auer, Degen, and Fischer (2013) as well as the citations therein.

[6] Bloom, Draca, and Van Reenen (2016) use plant-level data to argue that Chinese import competition led to greater research and development and patenting in their sample of European countries.

[7] Recent examples include Ashournia, Munch, and Nguyen (2014) and Utar and Ruiz (2013).

poor and middle-income developing countries. This paper examines the impact of rising Chinese export supplies on Mexico's local labor markets and relates to several strands in the literature on trade, development, and labor markets. In particular, a number of papers[8] studies Mexico's experience in the 1990s, including impacts on employment, wages, and inequality, after it began trade liberalization in the mid 1980s, culminating in the passage of the North American Free Trade Act.

This paper, instead, focuses on the post-NAFTA period in Mexico and the role that Chinese trade competition played in the evolution of Mexico's manufacturing industries. During this period, wage growth was smallest in the border states, which specialize in industries with the greatest overlap in Chinese export supplies. Border states also have among the highest mean wages in Mexico, and overall wage inequality declined during this period. I find that Chinese trade shocks led to consistently negative impacts across the conditional wage distribution, with workers in higher quantiles in manufacturing suffering slightly larger wage decreases than workers in lower quantiles. These striking results are among the first findings on the impact of Chinese trade on the wage distribution in an export-oriented, middle-income country.

It is possible to evaluate the impacts of trade, globalization, and outsourcing on multiple dimensions: at the firm, plant, regional, macro, and industry levels. This paper adopts the regional approach described in Autor et al. (2013), who provide a theoretically-grounded methodology to identify the causal impact of increasing Chinese exports on local labor markets. One advantage of the local labor markets approach is that it allows identification of spillover effects across industries. For instance, if trade shocks lead the least-skilled to lose their jobs in tradeable sectors and gain entry-level jobs in non-tradeable sectors, then a comparison between industries that are directly exposed to the trade shock and those that are not directly exposed would underestimate the impact of trade. Ebenstein, Harrison, McMillan, and Phillips (2014) examine the effects of trade and offshoring on wages in the U.S. and

[8] Goldberg and Pavcnik (2007) contains a detailed survey of many of these papers.

find no significant effects at the industry level, but they find substantial negative effects at the occupational level. The findings in this paper show that workers in manufacturing and outside of manufacturing show similar declines in wages in municipalities more exposed to Chinese export shocks.

While much of the focus in the recent literature, both theoretically and empirically, has been on how trade between low-income countries affects high-income counties, or vice versa, the effects of trade penetration from one developing country onto another are less studied. Mexico, like China, began a rapid shift from a fairly closed economy to an open one in the 1980s and is now firmly dependent on trade, especially with the U.S., its biggest trade partner. With China's rise, Mexico was arguably hit twice: first, via its domestic suppliers facing import competition from China, and second, via its customers in the U.S. potentially switching from Mexican to Chinese goods. Moreover, while other developing countries may also benefit from China's rise by being able to export minerals and raw materials to China (Hanson, 2010), Mexico exports relatively little of these commodities.

It is also important to consider institutional differences among high-income and low-income countries when considering how persistent labor market effects are likely to be. In standard economic theory we would expect for transitory trade shocks to be arbitraged away across space and for labor markets to adjust via mobility and firm reallocation in the long run. Jakubik and Kummritz (2017) argue that precisely this is happening in the U.S. On the other hand, Dix-Carneiro and Kovak (2017) find that the effects of tariff cut-induced shocks to labor markets in Brazil have grown, rather than being arbitraged away, over the course of twenty years at the regional level.

Mexico is also an interesting case study because its potential margins of adjustment are different. Mexico has no social safety net to deal with major labor demand shocks, and employment remains high even in crises, though workers are more likely to exit the formal sector and enter the informal sector during crises. Union protections are weak to nonexistent, no centralized wage bargaining mechanism exists for workers in the manufacturing industry, and hence workers have little bargaining power. It follows that wages are a potentially more important margin of adjustment here than in Europe or the U.S., where a wide body of work has found evidence of downward wage rigidities.

To study the effects of rising Chinese export supplies, I use cross-sectional data from the decennial Population Censuses in 2000 and 2010. Thus, the difference in local labor market outcomes between 2000 and 2010 identifies the medium-run impact of the Chinese shock. This period is ideally suited for the question because Chinese exports to Mexico, and to much of the rest of the world, started rising sharply in 2001, the year when China acceded to the WTO. The data allow me to control for a rich set of characteristics also important in the evolution of manufacturing employment and wages and provide information representative of the entire country using consistent municipality identifiers across time. Other work on Mexican labor markets has typically used the country's labor force survey, but the survey is not representative at the municipality level, did not cover much of the country prior to 2005, and its weighting was altered in 2005, making comparisons across time not feasible (Bosch and Campos-Vazquez, 2014).[9] Although the Population Censuses allow me to identify net impacts at the local labor market level, they do not, however, allow me to match workers to firms or to perform analysis at the plant level.

The identification strategy takes advantage of the fact that municipalities in Mexico are differentially affected by Chinese export shocks due to their industry structure prior to China's growth. Research using a similar methodology and focusing specifically on the role of rising Chinese trade typically finds negative impacts. Autor et al. (2013) find negative impacts on local wages, though these effects are entirely in non-tradeable sectors; they detect no impacts on wages in manufacturing.[10] Other work (Ashournia, Munch, and Nguyen, 2014; Donoso, Martín, and Minondo, 2015; Malgouyres, 2017) finds negative impacts in Denmark, Spain, and France, respectively. Balsvik, Jensen, and Salvanes (2015), however, do not find

[9] The survey is now called Encuesta Nacional de Ocupación y Empleo (The National Survey of Occupation and Employment). It has been altered and expanded several times since the 1970s, when it first began.

[10] In a companion piece exploiting longitudinal data on workers in the U.S., Autor, Dorn, Hanson, and Song (2014) find persistent, negative wage effects.

any impact of rising Chinese exports on wages in Norway.[11] These differing results point to the importance of local institutions and policies in establishing whether labor demand shocks cause adjustments along employment, wage, or labor mobility margins.

3.2 Chinese growth, manufacturing, and regional exposure in Mexico

3.2.1 Chinese export growth

The share of the population employed in manufacturing declined during the 2000s in Mexico, as in the U.S., although not in absolute numbers. Figure 1 shows the evolution of employment in manufacturing, services, agriculture, and other goods-producing sectors (mining and construction), relative to baseline levels in the 1999 Mexican Economic Survey.[12] Employment has increased sharply in services, but manufacturing employment has increased modestly as well, despite the combination of Chinese export shocks, skill-biased technological change, and other factors typically associated with declines in overall manufacturing employment in high-income countries in the last two decades.

When Mexico adopted an export-oriented strategy, policymakers hoped to emulate East Asian economies (Korea, Taiwan, Hong Kong) that had kickstarted their own development with simple export-assembly operations and gradually transitioned into producing more complex, higher-value products (Hanson, 2010; Enright, Scott, and Dodwell, 1997). Mexico's manufacturing industries, however, continue to be labor-intensive. Hanson (2010) suggests that Mexico's dependence on labor-intensive manufacturing, as well as the overlap between the industries that both countries have specialized in producing, made it particularly vulnerable to China's rise in the 2000s.

To get a sense of China's rise, Figure 2 shows the growth in the value of imports from

[11] They note that Norway's centralized wage bargaining probably induces effects along the employment, rather than wage, margin.
[12] These data are computed from the Mexican Economic Censuses from 1999, 2004, 2009, and 2014.

China into Mexico, the United States, as well as into other large economies in Latin America for comparison. All values are in U.S. dollars, adjusted for inflation, and then normalized to 100 in 1995. Even accounting for the large rise in imports into the U.S. over this period, the rate of growth of Chinese imports into Mexico has been far more dramatic: imports have increased close to a hundredfold over this period. This growth became especially dramatic after 2001, when China entered the World Trade Organization. In particular, Chinese trade made up less than 2% of the value of Mexico's imports in 2000 and surged to about 18% of imports in 2015 (see Figure 3). Figure 4 shows the growth of imports into the U.S. from China, Mexico, Canada, and the rest of the world. Growth from China and Mexico closely tracked each other until about 2001, when Chinese import growth surged ahead of other countries. Imports from Mexico continued to increase at a more modest rate over this period.

One can imagine a counterfactual exercise that explores what U.S. imports from Mexico would have looked like had China not increased its exports in the 2000s. Hanson and Robertson (2010) perform such an exercise using a gravity model of trade and find that Mexico—while more exposed than other countries in their sample—would have faced only moderately higher export demand had Chinese exports stayed constant. They conclude that rising Chinese exports have served as a modest negative shock to demand for Mexico's exports.

3.2.2 Industry structure and regional variation

The identification strategy of this paper relies on variation in exposure to Chinese exports across industries, interacted with the importance of these industries in each regional economy. Before moving on to the theoretical framework and the estimation, this section briefly describes industries and spatial patterns in manufacturing in Mexico.

China has not increased its exports uniformly across sectors. Using a two-digit SITC classification to illustrate the main points, China's biggest exports worldwide in 2010 were in office machines, electric machinery, and telecommunications. Other large industries, such as metalworking, paper and wood, and pharmaceuticals, have had less import exposure in partner countries. A similar pattern plays out when looking at exports to Mexico only. While Mexico specializes in some of these industries, it also produces passenger vehicles and food products, which benefit from having lower transportation costs to the U.S. market, and which have been less affected by direct import competition from China.[13]

Mexico's manufacturing industries are concentrated in the country's northern border region with the U.S., as well as its densely populated central belt, which includes major cities like Mexico City and Guadalajara. As in the U.S., manufacturing exhibits agglomeration economies. For instance, Mexico's shoe industry is concentrated around the city of León, Guanajuato. While smaller export-assembly plants (maquiladoras) in sectors like electronics, which were heavily exposed to Chinese trade competition since 2000, are more likely to be in the border region, other types of manufacturing are more prevalent in the interior of the country. For example, the food industry is clustered around the Mexico City metropolitan area, and the automotive industry has established plants throughout states in the interior.

This spatial concentration of industries provides a suggestive interpretation of wage changes between 2000 and 2010 in Mexico. As Figure 5 illustrates, while mean wages increased across all regions in Mexico, wage growth was smallest in the border region. The extent to which local labor markets more exposed to Chinese import competition suffered greater wage declines is investigated in the causal analysis that follows.

3.3 Empirical estimation of China's export growth

3.3.1 Conceptual framework for evaluating regional impacts of China's rise

This paper studies the impact on local labor markets of rising exports from China. The growing body of work taking a local labor markets approach is based on the theoretical

[13] See Mendoza Cota (2015) for a detailed analysis of Mexico's comparative advantage across manufacturing industries.

framework in Autor et al. (2013), which I describe in what follows. Each region inside of a country is modeled as a small open economy. The purpose of this framework is to evaluate how a rise in China exports affects each region's labor market outcomes, such as employment and wages, as well as other outcomes that may be tied to local labor markets, such as marriage, divorce, and human capital. Autor et al. (2013) link trade quantities, rather than trade prices, into outcomes in labor markets because data on industry import prices are not available, an approach taken in this paper as well.[14]

Each region produces goods, which can be tradeable (e.g. manufacturing) or non-tradeable. A positive productivity shock leads Chinese exports to rise. This may happen as the economy shifts from centralized planning to being market-based or as institutions, policies, and productive capabilities improve following WTO accession. Importantly, the productivity shock results from forces that are internal to China (e.g. not because of rising demand in an external country). The impacts on employment and wages depend on China's change in demand for that region's exports and change in demand for that region's output to any market in which it competes with China. It follows that as China increases its exports in a good produced by a region, wages and employment in the tradeables sector decrease, and employment in the non-tradeable sector increases. Of course, the opposite story holds if China's demand for imports rises, but as a practical matter this channel is less important in the case of Mexico: while China presently makes up about 18% of Mexico's imports (the second biggest source following the U.S.), China receives less than 2% of its exports (with the U.S. receiving over 80%). Chinese-Mexican trade is particularly imbalanced, which in the context of the theory implies that a reduction in labor demand in the tradeables sector as import penetration increases is not balanced by an increase in labor demand as export opportunities grow.

[14] There are alternative frameworks for describing the impact of Chinese growth on other countries. Kovak (2013) and Topalova (2010) focus on trade prices rather than quantities but also take a regional approach. Ashournia et al. (2014) describe a partial equalibrium trade model in which Chinese export growth alters firms' product demand and apply it to study firm-level outcomes.

3.3.2 Empirical approach

The conceptual framework described above shows how a shock to Chinese supply alters the demand for goods produced in each local labor market. A region will be more impacted by Chinese trade growth when Chinese exports in a particular sector increase and when the regional share of employment in that sector is greater. To illustrate what this means, suppose that Chinese exports of electronics to Mexico increase. Only places that produce electronics in Mexico are directly affected. If a particular region contains, say, a tenth of Mexico's total employment in electronics production, that region gets assigned a tenth of the rise in imports from China. This calculation is applied to each tradeable sector and then these import measures are summed to generate a measure of a region's import penetration in dollars. The calculation is then scaled by the total number of workers in the region to arrive at a measure of import penetration per worker.

Formally, I create the following measure of a municipality's import exposure to China:

$$\varDelta Exp_m = \sum_{j \in \mathbb{M}} \Big(\frac{L_{m,j,t=1998}}{L_{j,t=1998}} \frac{\varDelta Imports_j}{L_{m,t=1998}} \Big). \tag{3.1}$$

In this expression, $L_{m,j,t=1998}$ denotes employment in industry j (where $\mathbb{M}$ refers to the set of industries in manufacturing) in municipality m in 1998, $L_{j,t=1998}$ denotes aggregate employment in Mexico in manufacturing industry j in 1998, $\varDelta Imports_j$ denotes the change in Mexican imports from China between 2000 and 2010, and $L_{m,t=1998}$ denotes total employment in municipality m in 1998. The first fraction represents the share of demand for a given industry that local manufacturers produce. The change in import values represents the increase in China's export supplies. The measure is divided by the total number of workers in the region to proxy for local labor market exposure at the worker level. Figure 6 illustrates import exposure across Mexico: it is highly correlated with manufacturing, with places mainly in the north and central region most exposed.[15]

[15] Large municipalities with the highest exposure include Nogales, Agua Prieta, and San Luis Río Colorado in Sonora; Juárez and Chihuahua in Chihuahua; Matamoros in Tamaulipas; Acuña in Coahuila; and Tecate

As noted in Autor et al. (2013) and subsequent work, the identifying variation in this measure depends on the degree of the employment mix in manufacturing, as well as how specialized municipalities are in industries that are subject to increases in export supplies. Since the resulting measure is correlated with how much of the local employment base is composed of manufacturing, I control directly for the manufacturing share of employment in the econometric specifications. The remaining variation, then, is based only on local concentration in industries subject to import shocks after adjusting for local manufacturing concentration.

To identify the relationship between import exposure and changes in local labor market conditions, I estimate variants of the following specification:

$$\mathit{\Delta} Y_m = \gamma + \beta \mathit{\Delta} Exp_m + \boldsymbol{X}_m^{'} \delta + \boldsymbol{\lambda}_r + \varepsilon_m. \tag{3.2}$$

$\mathit{\Delta} Y_m$ denotes the change between 2000 and 2010 in the outcome of interest (e.g. the average of the natural logarithm of wages or the share of the working-age population employed in manufacturing in municipality m). Some specifications also contain a vector, denoted here by $\boldsymbol{X}_m$, that contains a rich set of variables controlling for baseline characteristics of the labor force and demographic composition that may be related to changes in manufacturing employment. The vector $\boldsymbol{\lambda}_r$ contains dummies for five of six regions in Mexico. These regional dummies flexibly control for region-specific trends in outcomes in a first-differenced specification. The coefficient of interest is β, which indicates the effect of a change in a municipality's import exposure to growing Chinese exports. Standard errors are clustered at the state level to control for spatial correlation.

A potential issue with the estimation strategy here is the endogeneity of the import exposure measure since an unobserved domestic demand shock could be picked up by the import measure. For instance, an industry facing a positive demand shock may grow domestically

in Baja California. All of these are in border states and are notable for their maquiladoras. El Salto in Jalisco has the highest import penetration among municipalities deeper in the interior of the country.

and at the same time import more from China. In such a scenario, the OLS estimator of β would be biased up (if the true effect is negative). In general, industry-based shocks, either from domestic demand or from foreign trade in other countries, could also affect Mexican imports from China, leading to a positive or negative bias, depending on the nature of the shock.

Following Autor et al. (2013) and other researchers studying the impact of an export supply shock in China, I instrument domestic changes in imports with changes in imports into other countries also affected by the supply shock. Formally, the instrumental variable is defined as follows:

$$\Delta Exp_m^o = \sum_{j \in \mathbb{M}} \Big(\frac{L_{m,j,t=1998}}{L_{j,t=1998}} \frac{\Delta Imports_j^o}{L_{m,t=1998}} \Big). \tag{3.3}$$

This expression replaces imports into Mexico with the sum of exports into a basket of similar countries, denoted by $\Delta Imports_j^o$. This instrumental variables strategy relies on the assumption that the main driver in the import measure is a positive shock to Chinese productivity, which would have increased Chinese exports across a similar basket of goods in other countries. This is reflected in the similarity across countries in which goods are imported from China and in the well-documented history of factors that led to China's manufacturing growth, as described in a previous section.

Threats to identification can come from several sources, which also arise in the case of Autor et al. (2013). Demand shocks within sectors can be correlated across different countries. For instance, the U.S. and Mexico share economic cycles and are major trading partners. In the analysis, I use a set of countries (Bolivia, Chile, Colombia, Spain, and Venezuela) that are comparable to Mexico and likely to face similar supply shocks from China, but that are not among the top exporters to or importers from Mexico.[16] Robustness checks using other middle-income countries with available data do not alter the results in

[16] Mexico's top export destinations are the U.S. and Canada. All other countries make up less than 2% of Mexico's exports. Top exporting countries to Mexico are (in descending order) the U.S., China, Japan, Germany, and Korea.

this paper. It is also theoretically possible that some sectors in Mexico suffer negative productivity or technology shocks, leading to a rise in Chinese imports from these sectors. Although I cannot directly refute this, the available evidence suggests that the sudden and dramatic growth in Chinese productivity has been responsible for China's increasing share of exports, rather than processes internal to Mexico.

Finally, it is important to note that the measure, as constructed here, is designed to serve as a proxy for a local labor market's exposure to rising Chinese trade supplies, but it is possible that the measure in Equation (1) misses two important channels by which Chinese exports can affect local employment and wages. The first channel is that China may displace some of Mexico's export capacity, and this measure misses the impact of the potential loss in Mexican exports. I address this issue in a later section by estimating the total local effect from domestic exposure to rising Chinese export supplies into Mexico plus international exposure to rising Chinese trade with the U.S.

The second channel that this measure may not account for includes certain kinds of general equilibrium effects. Using shoes as an example, suppose that Mexico imports only shoes from China, but not the intermediate goods involved (e.g. leather, rubber). In this case, the measure captures the direct effect on producers that make shoes, but it does not capture the indirect effect on producers that make leather and rubber. Presumably, if shoe producers in Mexico are adversely impacted by increasing competition from China, then domestic producers of leather and rubber may be adversely impacted as well. If there are strong agglomeration economies, such that final and intermediate good producers are located in the same labor market, then the measure here still captures a net effect on local labor markets. If the supply chain is more spatially dispersed, and (continuing the same example) if there is a negative effect on both the shoe producers and the leather and rubber producers, but only domestic imports of shoes increase, then the coefficient on import exposure would be attenuated. I do not pursue this approach here, but one way of dealing with this channel would be to construct a regional input-output table using Mexico's national input-output table. Such an approach would capture the effects on local areas via input-output linkages and allow for the identification of the indirect effect of Chinese imports on sectors in Mexico. The indirect effect could then be added to the direct effect for each local labor market to arrive at a measure of import exposure that could be used in the econometric specifications here.

3.3.3 Data and measurement

Data on wages, aggregate employment outcomes, and human capital come from the Mexican Population Censuses in 2000 and 2010. Information from the 2000 Population Census is used to provide the start-of-the-period covariates in Equation (2). I use the municipality as the unit defining a local labor market. In some cases, municipalities have split over time, and urban agglomerations may sprawl across municipalities. I use identifiers for Mexican municipalities that use consistent boundaries over time.

I use trade data from the UN Comtrade Database on Mexican imports at the four-digit SITC (Rev. 2) level. Data on employment by sector in Mexico comes from the Mexican Economic Census, which is conducted every five years. The last census conducted prior to China's entry into the WTO was in 1998. Using slightly lagged measures of employment reduces concerns with firms adjusting business operations in anticipation of China's entry into the WTO and potentially introducing endogeneity into the import penetration measure. Since the Economic Census uses the Mexican version of the North American Industry Classification System, rather than the SITC or HC mappings, I build a correspondence table between SITC and NAICS codes, based in part on the table provided by Feenstra and Lipsey and filling in missing values where necessary.[17] Trade values are converted from U.S. dollars to Mexican pesos and deflated using the Mexican CPI.

[17] Their mapping is provided at http://www.nber.org/lipsey/sitc22naics97/.

3.3.4 Impacts on local labor markets

To evaluate the impact of Chinese import penetration on the local change in the share of the working-age population employed in manufacturing, I estimate Equation (2) without any covariates aside from the measure of Chinese import penetration defined in Equation (1). Column (1) in Table 3 indicates that every unit increase in import penetration (that is, a 1,000 peso increase) led to a decline of 0.153 percentage points in the share of the working-age population employed in manufacturing. This coefficient is statistically significant from zero at the 1% level. Since the mean of the difference in import exposure is about 11.6, this indicates that an increase from no Chinese penetration to the average would have led to a 1.8 percentage point decline in the share of the working-age population employed in manufacturing, which is slightly below the mean decline observed over the ten-year period in the data.

Column (2) includes the manufacturing share of employment in 2000 as a regressor. In general, import penetration is correlated with the manufacturing share, and manufacturing was slowly declining over this period. Hence, omitting manufacturing share may lead the import penetration measure to pick up a negative trend in manufacturing. Indeed, the coefficient on manufacturing share is negative, indicating that places with more manufacturing employment lost relatively more manufacturing jobs, and the coefficient on the import penetration measure becomes much smaller in magnitude. This indicates that once variation in baseline levels of manufacturing employment is controlled for, variation across how exposed different sectors were in a municipality to Chinese export supply shocks does not have a major impact on changes in the manufacturing share. Column (3) includes an additional set of covariates, controlling for a rich set of baseline characteristics and regional trends that may be associated with changes in manufacturing employment, to isolate the impact of a change in import penetration between 2010 and 2000. These controls include the male share of the working-age population, the share of individuals into each of four mutually exclusive groups based on educational attainment,[18] the share of the population in agricultural employment, and the female employment share. Once these explanatory variables are added, the coefficient on import exposure is now very close to zero.

One possibility, raised earlier, is that the OLS estimator is biased, perhaps because local demand shocks lead to increases in employment in certain sectors and increases in imports from China at the same time. This would attenuate any negative effect from Chinese import penetration. I use 2SLS estimation to address this issue. Column (4) shows the first stage, where the import measure defined in Equation (1) is regressed on the measure in Equation (3), as well as all of the covariates in the previous column. The relationship is precisely estimated and strong, with a Kleibergen-Paap F statistic of 135. The resulting instrumental variables estimate, shown in Column (5), remains small and not statistically significant from zero.

Although these results may appear surprising since Autor et al. (2013) find negative impacts on manufacturing share in the U.S., subsequent work in other contexts has not consistently found similarly large negative effects. Balsvik, Jensen, and Salvanes (2015) find negative effects in Norway, but they are much smaller in scope. They estimate that the Chinese import shock explains less than 10% of the observed decline in manufacturing in Norway. In contrast, Taniguchi (2017) finds that Chinese export supplies *increased* manufacturing employment in Japan.

To provide a richer perspective on local impacts and identify effects aside from the extensive margin of manufacturing employment, I proceed with 2SLS estimation of other labor market-related outcomes using the strategy outlined above. Table 4 shows the results of 2SLS estimation of Equation (2) on other labor market outcomes. All specifications include the full set of explanatory variables in Column (3) in Table 3. The results show that the proportion of individuals with at least secondary schooling and the mean years of schooling among

[18] These groups include individuals with than primary schooling completed, individuals with primary schooling, individuals with at most lower secondary schooling completed, and individuals with at least secondary schooling. Having less than primary schooling is the omitted category.

the working-age population both decline [Columns (1) and (2), respectively], indicating that import penetration leads to a reduction in human capital in affected municipalities. Column (3) indicates that a 1,000 peso increase in Chinese exports also reduces the proportion of high-skilled individuals in manufacturing by 0.193 percentage points.[19] In Columns (4) and (5) I evaluate the impact of import exposure on the proportion of the working-age population that is employed and unemployed, and find a small, imprecisely estimated negative impact on the former, and a small, imprecisely estimated positive impact on the latter.[20]

These results provide evidence that even if the overall manufacturing share of the working-age population is unaffected, the composition of employment in manufacturing, as well as the composition of the labor force, changes. We see, then, that the effects of Chinese import penetration are muted once a rich set of explanatory variables is included, pointing to other drivers as being more important in the reduction of manufacturing employment in Mexico. Nevertheless, effects on the intensive margin appear to play an important role in altering the structure of employment in response to the Chinese trade shock.

3.3.5 Wage effects

Table 5 shows the results of estimating Equation (2) on mean log wages at the local level, with the outcome variable expressed as the difference in log points.[21] The first three columns show OLS results as more explanatory variables are added, and Column (4) shows instrumental variables results with the full set of covariates. Estimates of the effect of import exposure are negative and statistically significant from zero in all specifications. The size of the effect declines as controls are added, but it remains sizable even with a rich set of controls. The instrumental variables estimate indicates that a 1,000 peso increase in import exposure

[19] I define high-skilled as having attained secondary schooling or more, as college education remains relatively rare in Mexico.

[20] I also find a positive impact on the log of the overall population that is unemployed. On the other hand, I find negative impacts on the overall population, employment, and employment in manufacturing, also measured in logs.

[21] I define wages as monthly earnings divided by (4.3 x hours worked per week) for workers with at least 30 hours of work per week. Wages are trimmed at the top and bottom by 2% to deal with topcoding and zeroes. Results are not sensitive to this specification.

reduces wages by 0.536 log points. This translates into approximately a 5% relative decline in wages for a municipality going from the 25th percentile in exposure to the 75th percentile. That the 2SLS estimate is larger in magnitude than the OLS estimate is evidence of some combination of measurement error and endogeneity, i.e. of strong local demand conditions possibly inducing greater imports, which would lead to an upward bias in the estimated impact of an increase in Chinese exports. Also note that since wages increased over this period, a negative coefficient here amounts to smaller wage growth in municipalities more exposed to the trade shock.

I proceed to decompose the wage effect by focusing on subgroups. Table 6 shows the 2SLS results. The effects on men and women are shown in Columns (2) and (3). The results are remarkably similar. Columns (4) and (5) focus on manufacturing and non-manufacturing and indicate that wage losses are substantially greater in the former.[22] Since I do not have longitudinal data on workers' wages, I cannot establish directly whether these changes are reflected as relative decreases for the same workers or whether differential selection into or out of manufacturing is leading to these results. However, when combining these results with the earlier findings on overall employment, human capital, and skill levels within manufacturing, the evidence indicates that more-skilled workers in manufacturing are replaced with less-skilled workers. I return to this point when I consider impacts on the conditional wage distribution in the next section.

3.3.6 Impacts on the conditional wage distribution

A set of studies on inequality in Mexico from the 1990s and early 2000s argued that trade liberalization increased inequality.[23] However, other research has noted that i) the effects of earlier rounds of trade liberalization may have differed from the effects of the passage of

[22] I also look at differences by age group and find consistent results across age by gender cells (results not shown).

[23] A discussion and citations of this work can be found in Goldberg and Pavcnik (2007).

NAFTA (Robertson, 2004)[24]; ii) Mexico experienced other concurrent policy shocks, such as labor market reforms, deregulation, declines in the minimum wage (Bosch and Manacorda, 2010), and so on; iii) the effects of NAFTA may have already been priced in prior to its actual passage; iv) regions across Mexico were differentially exposed to trade, potentially biasing estimates that did not take this into account (Chiquiar, 2008); v) earlier studies did not give enough time for Stolper-Samuelson effects to materialize (i.e. an application of the point made in Slaughter, 2000); and vi) tariff cuts were not evenly distributed across sectors, but rather affected low-skilled workers the most since they were employed in the most protectionist sectors prior to trade liberalization (Hanson and Harrison, 1999).

More recent studies, however, document that wage inequality declined in Mexico after NAFTA, i.e. during the period of study in this paper (Campos, 2013; Lustig et al., 2013) and provide evidence in support of supply and demand factors for skilled versus unskilled labor, but do not consider the role that trade might have played. Figure 7 illustrates the change in mean log wages across Mexico for each quantile in the wage distribution. It is clear that the greatest gains have accrued to those in the bottom of the wage distribution. A similar pattern holds if controlling for region.[25]

It is well-known that trade shocks can impact the wage distribution, depending on the incidence of labor demand shocks. The results thus far indicate that mean wages decline and the skill content of manufacturing declines. Chetverikov, Larsen, and Palmer (2016) propose an asymptotically unbiased estimator, which they refer to as grouped IV quantile regression, that allows identification of the distributional effects of an endogenous treatment (i.e. a trade shock) when the treatment varies at the group (i.e. local labor market) level.[26]

The estimator is applied as follows. First, I calculate the quantile q of the natural

[24] To be precise, the trade liberalization of the 1980s led to Mexico's opening to trade with other low-skill labor abundant countries, whereas the later signing of NAFTA led to decreasing trade barriers to high-skill abundant countries.

[25] The only exception here is the north of the country, i.e. the group of states south of the states bordering the U.S., where the effects are evenly distributed across wage quantiles.

[26] The reason for adopting this methodology is that conventional quantile regression techniques are inconsistent in the presence of group-level unobservables.

logarithm of wages in each municipality using data from the decennial Population Censuses. Then I replace the difference of the average outcome in Equation (2) with the difference of the respective quantile, ΔY_m^q. I also continue to instrument ΔExp_m with ΔExp_m^o. The results for the full wage distribution are shown in Figure 8. The effect appears fairly constant, negative, and precisely estimated throughout the wage distribution. Splitting the sample into men and women (see Figures 9 and 10), I find constant effects for men and slightly more negative effects for women at the upper end of the wage distribution. When I separate the sample into individuals employed in manufacturing and those outside of manufacturing (see Figures 11 and 12), I find progressively more negative impacts at the upper end of the wage distribution for workers in manufacturing. Nevertheless, I cannot reject the hypothesis of constant effects along the wage distribution.

These results contrast with recent empirical studies on the U.S. and Europe, which typically show that the least-skilled see the biggest reductions in wages and employment prospects in response to Chinese export shocks.[27] It appears that a greater overlap between the skill distribution of the importing country and China results in more evenly distributed negative impacts on wages.[28] These results may explain a part of the puzzle in Campos (2013), who finds that, after increasing during a period of trade liberalization, wage inequality has declined in Mexico, in part due to a reduction in the return to high-skilled labor. The findings here indicate that exposure to Chinese trade penetration may have had a slightly

[27] Building on the insights in Melitz (2003), a large literature on firm heterogeneity analyzes firm survival and exit, product upgrading, and skill content in response to trade shocks. This work highlights the importance of innovation, skill upgrading, offshoring and vertical integration (e.g. Bloom, Draca, and Van Reenen, 2016; Bernard, Redding, and Schott, 2011; Grossman and Rossi-Hansberg, 2008; Costinot, Oldenski, and Rauch, 2011). Also evaluating the impact of Chinese exports, Mion and Zhu (2013) find that firms with greater exposure to Chinese export shocks increase wages and employment for non-production workers in Belgium. Álvarez and Claro (2009) also find negative effects on plant survival in Chile, but unlike much of the literature focusing on the U.S. and Europe, they find no evidence that plants adjust to rising Chinese export supplies by quality or skill upgrading. Instead, they find that plants shift to using more labor-intensive techniques and producing less complex products.

[28] Mion and Zhu (2013), Rodrik (2006), and Schott (2008) note that Chinese exports are more sophisticated than exports from other low-income countries. Relatedly, Chinese manufacturing wages were lower than wages in Mexico in 2000, but as of today the situation is reversed. Thus, it is difficult to make a clear prediction regarding Stolper-Samuelson effects as in the case of a high-income country opening to trade to a low-income country.

greater negative impact on workers in the upper end of the distribution in manufacturing, who are more likely to be high-skilled. This is also consistent with the pattern of inequality shown earlier between the Decennial Censuses.

3.4 Robustness of the results

3.4.1 Impacts on wages

To check the robustness of these results on wages, I perform several checks on the data. First, I check whether the largest city in Mexico is driving the results. I re-estimate Equation (2) omitting Mexico City; the resulting coefficient is shown in Column 2 in Table 7. It is virtually unchanged from the coefficient on the full sample, which is in Column 1. Since Mexico City sprawls deeply into neighboring areas, I also re-estimate the main specification by dropping the city itself, as well as the surrounding states of México and Morelos from the analysis. The resulting estimate, in Column 3 of Table 7, is slightly greater in magnitude. I then check the sensitivity of the results to outliers and drop municipalities in the bottom and top 3% of import exposure to China. Again, the coefficient increases in magnitude, but stays statistically significant from zero at the 5% level.

A potential concern with evaluating outcomes between 2000 and 2010 is the financial crisis of 2008, which spilled over into Mexico by 2009. As Figure 1 indicates, Mexico's imports from China decreased in 2009, as the recession dampened economic activity, although imports abruptly bounced back in 2010 and then continued increasing. I do not have reason to believe that municipalities disproportionately affected by the crisis were also disproportionately exposed to Chinese trade exports, but to verify if this is the case, I incorporate data from the 2015 Population Census. By this point, Mexico had recovered from the recession, so extending the analysis to a longer time difference allows me to evaluate whether the financial crisis is leading me to find spuriously negative wage effects in 2010.

A drawback of using the 2015 Population Census is that it does not have the rich set of information present in the Decennial Censuses. In particular, it has no measure of hours worked, so I cannot create a wage variable for full-time workers. Since monthly earnings are available in 2000, 2010, and 2015, I replace wages with earnings in the main specification. First, I use the hourly data available in the decennial Censuses to estimate the impact of the Chinese trade shock on earnings for full-time workers only. The estimate, shown in Column (5) of Table 7, is close to the estimate using wages, which implies that a negative shock to wages is driving the decrease in earnings, not a reduction in hours worked.[29] In Column (6), I extend the sample to show results for earnings for individuals with positive earnings. The full-time distinction makes little difference. Finally, I compare effects across the ten-year difference versus the fifteen-year difference in the data. The estimate in Column (7) drops from -0.590 to -0.372, indicating a smaller, though still economically significant, effect across the longer timespan.

In theory, one would expect labor mobility to eliminate differential labor market-related outcomes across municipalities in the long run (Blanchard and Katz, 1992; Bound and Holzer, 2000). The smaller effect of import exposure using the longer time difference supports this view. Note, however, that the Chinese export shock was not a discrete event, but rather a gradually increasing shock over time. In this sense, the comparison between results using the 2010 and 2015 Censuses is not a straightforward analysis of shorter versus longer horizons but a more complex function of how the shock has evolved (Dix-Carneiro and Kovak, 2017; Autor, Dorn, Hanson, and Song, 2014). Regardless, that earnings were substantially lower in 2015 in municipalities more exposed to Chinese export supplies indicates that labor market adjustments were still taking place.

3.4.2 Including the effects of export competition

A potential problem with the measure in Equation (1) in this paper is that it identifies the effect of rising Chinese exports solely through domestic exposure to imports. This may

[29] I also look directly at the impact of import penetration on hours worked and find no effect.

not be reasonable in the case of an export-oriented economy like Mexico. In an extreme case, it is possible that Mexico may not even experience an increase in imports from China in some industries. In that case, the measure used here would not detect a change in local import exposure. Suppose, however, that Mexico would have exported the output of these industries to its trading partners in the absence of growth in Chinese exports. This potential loss in exports is not accounted for in Equation (1).

To capture the effect of Chinese growth potentially displacing Mexican firms that produce for the domestic market plus the effect on firms that potentially lose exports to the U.S., I modify the measure in Equation (1) to create

$$\mathit{\Delta Exp}_m = \sum_{j \in \mathbb{M}} \Big(\frac{L_{m,j,t=1998}}{L_{j,t=1998}} \frac{\mathit{\Delta Imports}_{Mex,j} + \frac{Imp_{Mex,j}}{Imp_{Total,j}} \mathit{\Delta Imports}_{US,j}}{L_{m,t=1998}} \Big). \tag{3.4}$$

A similar calculation is performed in Autor et al. (2013). The difference from Equation (1) is in the second summation term. This term weights the change in imports into the U.S. from China by the baseline share of total imports into the U.S. coming from Mexico in each industry. This is added to the change in imports into Mexico from China, as done originally. The measure here estimates the total local effect from domestic exposure to rising Chinese export supplies into Mexico plus international exposure to rising Chinese trade with the U.S.

Once I include the local exposure to potential losses in the U.S. market, which accounts for over 80% of Mexico's exports, the mean change in import exposure rises by about 60%. When re-estimating Equation (2),[30] I find that the impact on manufacturing employment, once accounting for other important covariates, remains close to zero, as originally estimated. The magnitudes of the coefficients on wages, shown in Table 8, decline by about half or a bit less, which is consistent with the measure of import exposure now being scaled up (as in Autor et al., 2013). The trend in estimates and levels of statistical significance remain similar to what was found previously.

[30] I calculate the instrument in the same manner as before using the new measure of total import exposure in Equation (4).

One interpretation of these results is that the majority of the Chinese trade shock is channeled through its effects on domestic import exposure and not through the displacement of exports to other countries. The latter point is consistent with the results in Hanson and Robertson (2010), who estimate that Mexico's manufacturing export demand would have been either 0.1% or 0.2% higher (depending on the value of the import price index used) had China's export supplies remained constant in the period between 1995 to 2005. That is a small figure, indicating that China's growth has not had a large impact on Mexico's exports in the period immediately after China joined the WTO and its export growth surged. However, returning to the interpretation of the results, it is important to note that the change in imports from China into the U.S. across industries is highly correlated with the change in imports into the Mexico. Collecting data on the share of local output that is exported by industry would be helpful in addressing this point, such that one could identify whether local areas that export the output in a given industry are differentially affected from areas whose output is sold domestically.

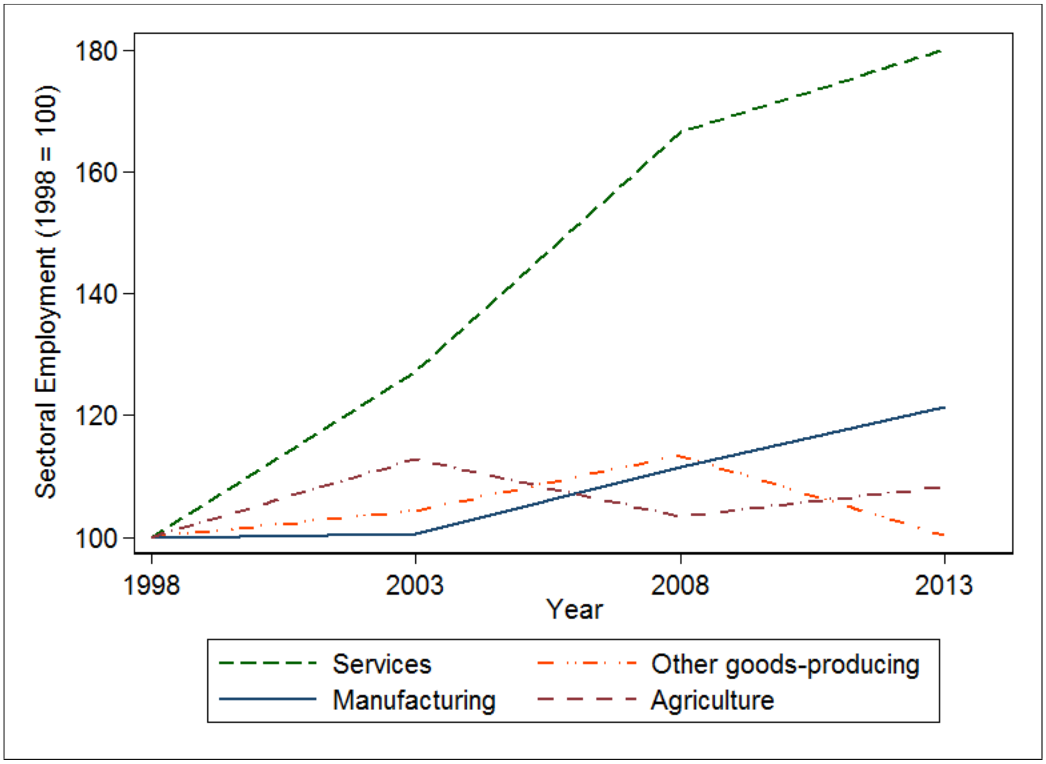
3.5 Conclusion

The so-called "China shock" arguably has been greater in Mexico than in wealthy countries. Returning to the question posed by Hanson (2010)—"Why isn't Mexico rich?"—I evaluate the role that the huge surge in Chinese export supplies has played on Mexico's labor markets after China's accession to the WTO. Contrary to studies in the U.S. and Europe, I find no negative impacts on manufacturing employment once I control for a rich set of covariates also affecting the evolution of manufacturing employment. On the other hand, I find that local labor markets exposed to greater Chinese trade, conditional on baseline characteristics and regional trends, have lower wages, show reductions in average schooling, and produce manufacturing with a less-skilled workforce. Results from quantile IV regressions indicate heterogeneous effects across the wage distribution in manufacturing, with individuals in the upper deciles showing slightly more pronounced negative effects.

These findings indicate that the local response to Chinese trade shocks differs across developing and wealthy countries. Firms in wealthy countries improve the quality of their products and hire a more-skilled workforce, indicating that there may be complementarities between workers in low-income nations and high-skill workers in developed countries. Firms in developing countries, where low-quality institutions, distorted incentives, and a less-skilled workforce likely produce barriers to climbing the product quality chain, may instead respond by cutting wages and becoming more labor-intensive to compete effectively.

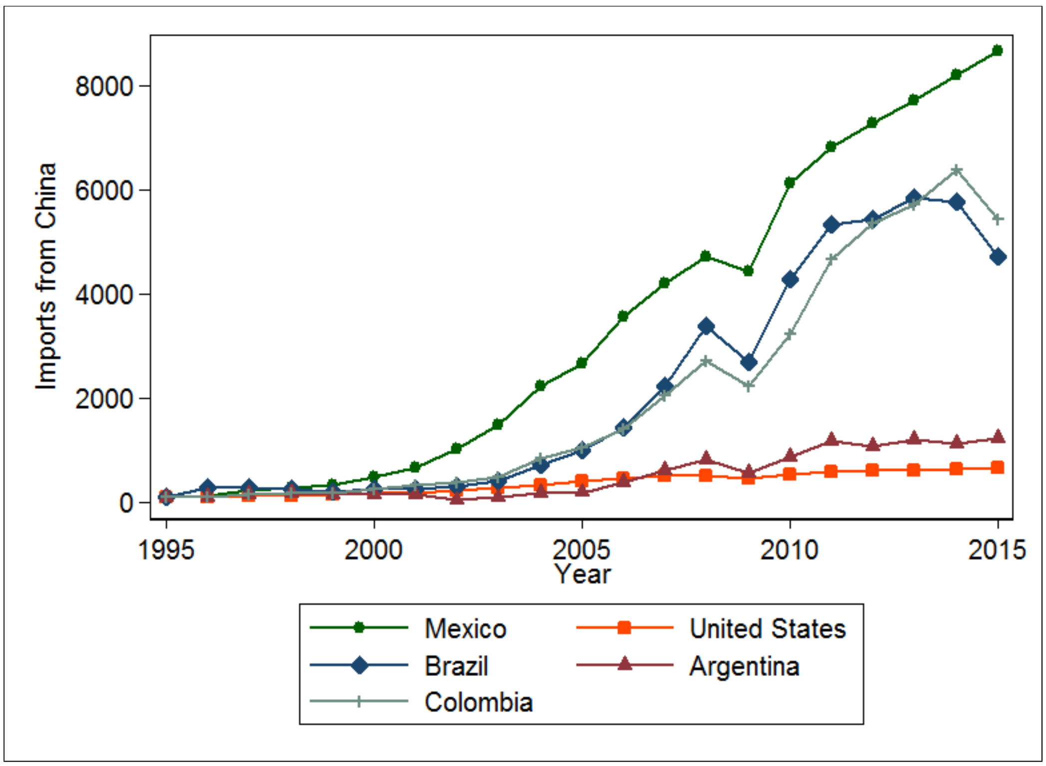
3.6 Figures

Figure 3.1: Growth in Sectors



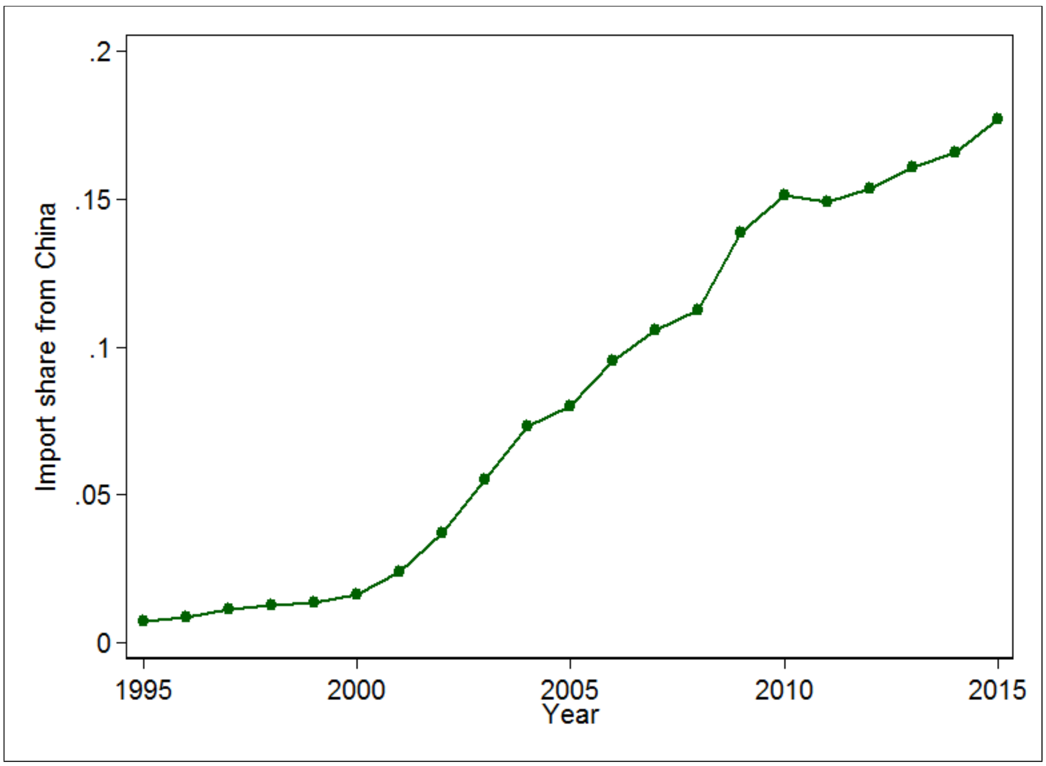
This figure shows the growth in employment in manufacturing, services, agriculture, and other goods-producing sectors between in Mexico, using data from the 1999, 2004, 2009, and 2014 Economic Surveys. All levels are normalized to 100 in the 1999 Survey.

Figure 3.2: Growth in the Value of Imports from China over Time



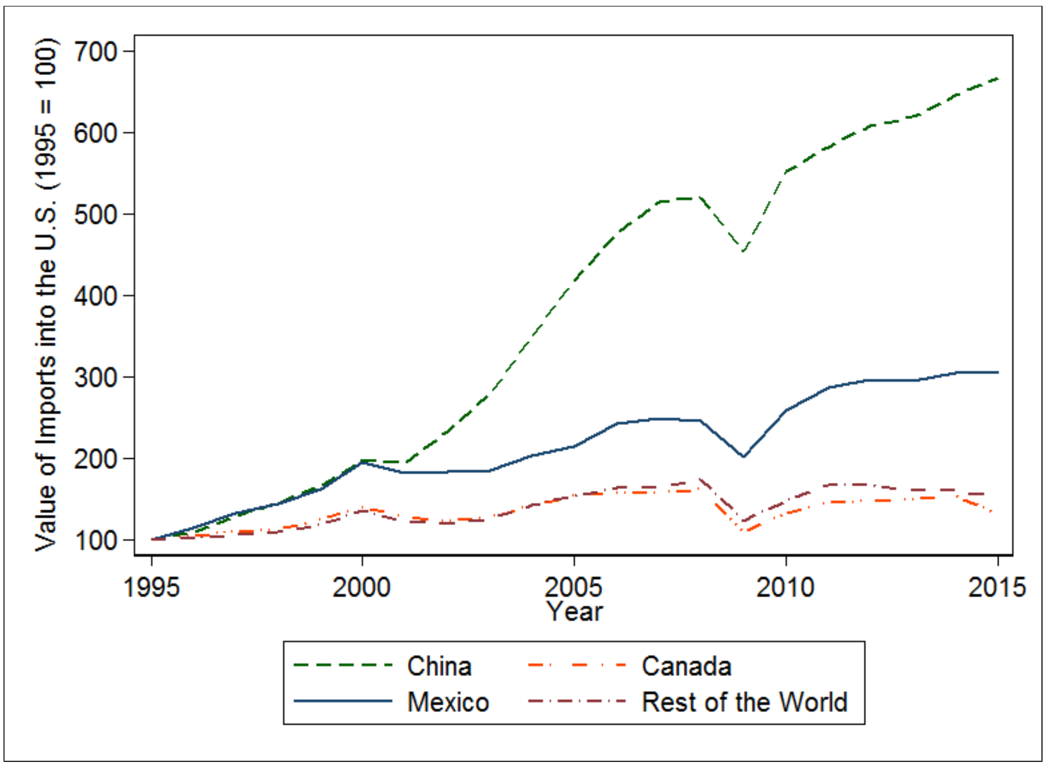
This figure plots import trade values from China into Mexico, the United States, Brazil, Argentina, and Colombia since 1995. All values are adjusted so that numbers in 1995 start at 100 to show relative trends in growth. Trade values are in 2010 dollars. Data are calculated using the information provided by the UN Comtrade Database and adjusted for inflation.

Figure 3.3: Growth in the Share of Imports from China over Time



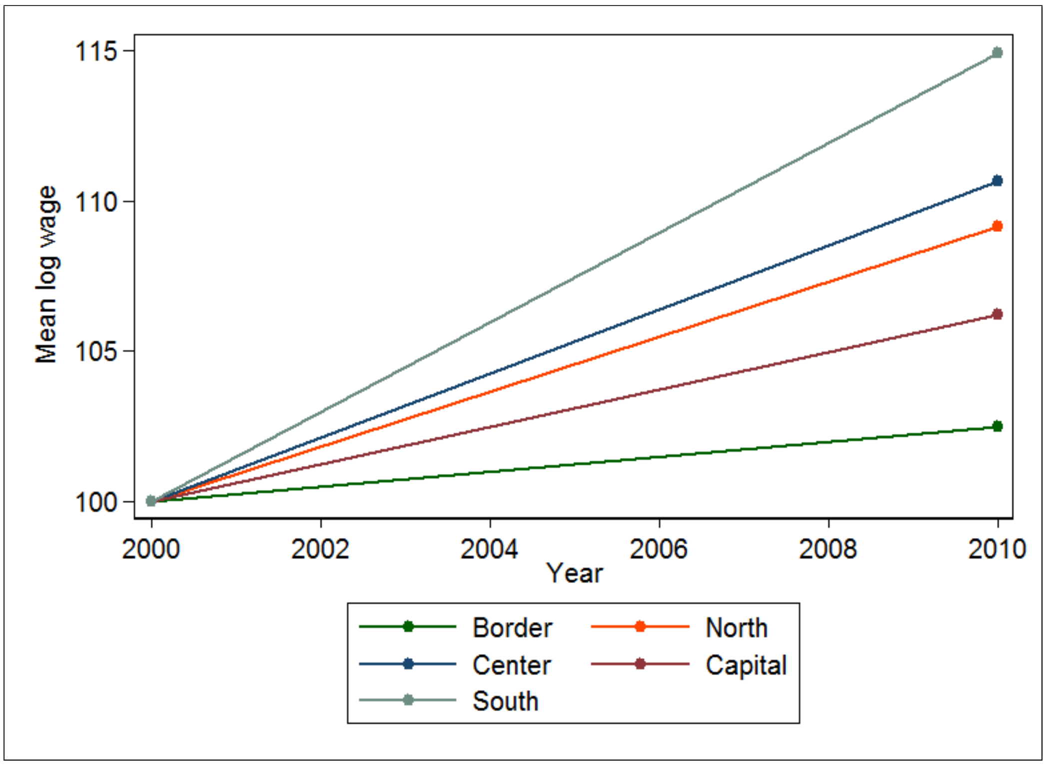
This figure plots import trade values from China into Mexico as a fraction of all imports. Data are calculated using the information provided by the UN Comtrade Database and adjusted for inflation.

Figure 3.4: Growth in Imports into the U.S. over Time



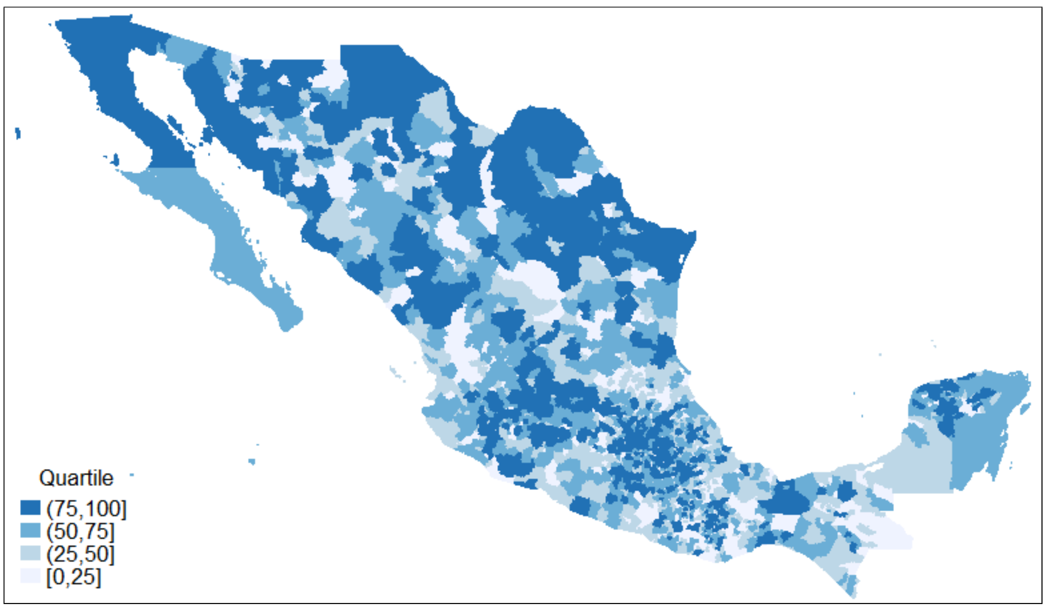
This figure plots import trade values from China, Mexico, Canada, and the rest of the world into the U.S. Data are calculated using the information provided by the UN Comtrade Database and adjusted for inflation. Values are normalized to 100 in 1995.

Figure 3.5: Wage growth across Mexican Regions



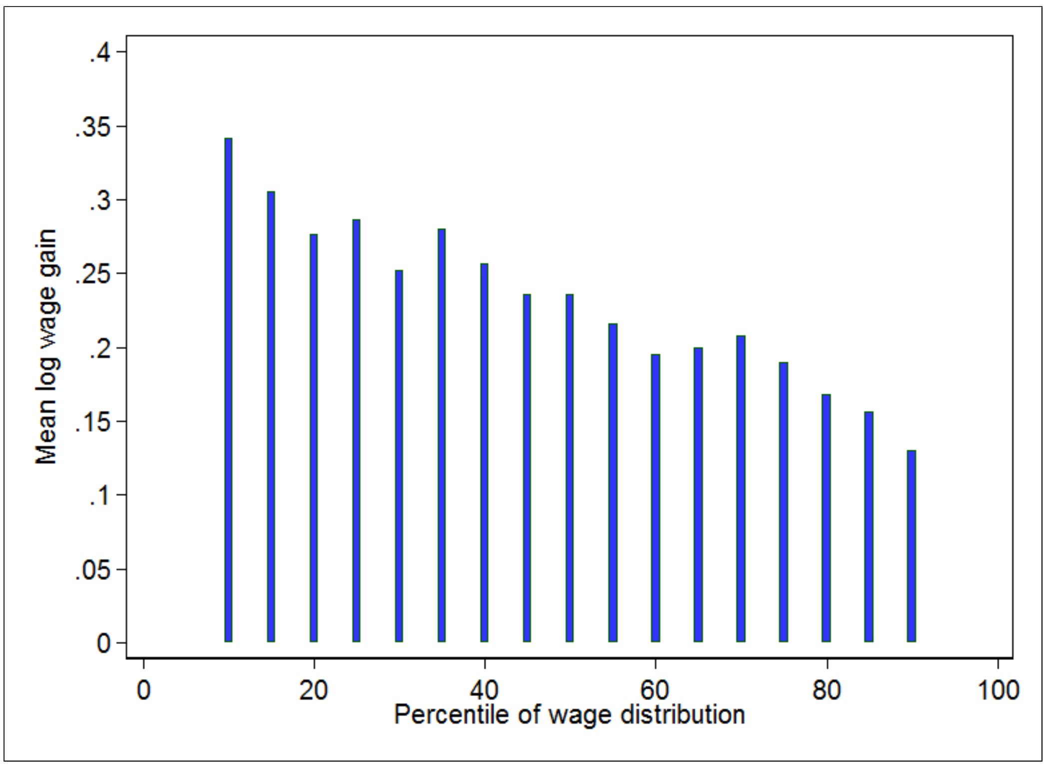
This figure shows the growth in mean (log) wages across Mexican regions between the 2000 and 2010 Population Censuses. All wages are normalized to 100 in 2000.

Figure 3.6: Import exposure across Mexico



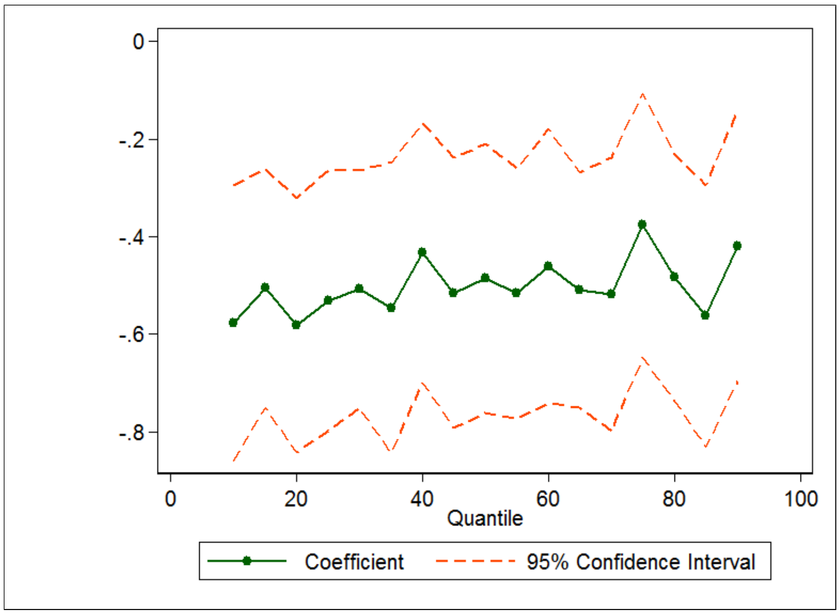
This figure shows the difference in import exposure between 2000 and 2010 across Mexico by quartiles. All municipalities are ranked by their exposure, from no exposure (0th quartile) to most exposed (100th quartile).

Figure 3.7: Differences in Mean Log Wages across the Wage Distribution



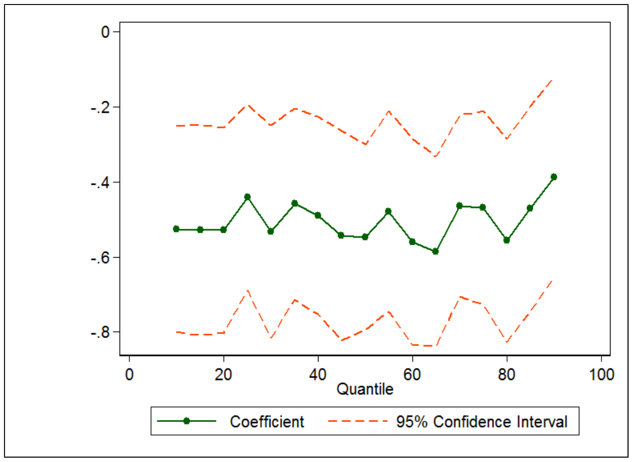
This figure shows the change in mean real log wages between 2000 and 2010 in Mexico across the wage distribution. Data come from the decennial Population Censuses.

Figure 3.8: The Impact of Chinese Exports on the Conditional Wage Distribution on the Full Working-age Population



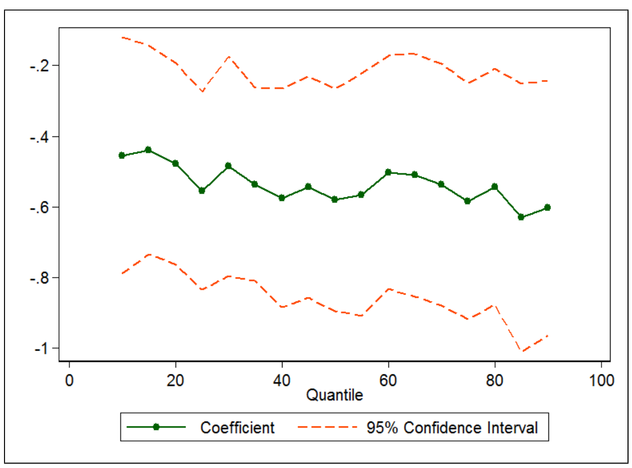
This graph incorporates the full working-age sample. The figure plots grouped instrumental variables quantile regression estimates of the effect of a 1,000 peso increase in Chinese imports per worker. Point estimates and confidence intervals for each quantile are shown. Units on the vertical axis are log points.

Figure 3.9: The Impact of Chinese Exports on the Conditional Wage Distribution on Men



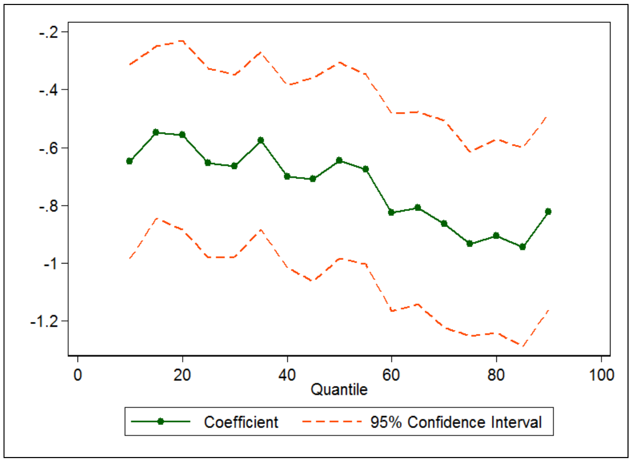
This graph incorporates working-age men only. The figure plots grouped instrumental variables quantile regression estimates of the effect of a 1,000 peso increase in Chinese imports per worker. Point estimates and confidence intervals for each quantile are shown. Units on the vertical axis are log points.

Figure 3.10: The Impact of Chinese Exports on the Conditional Wage Distribution on Women



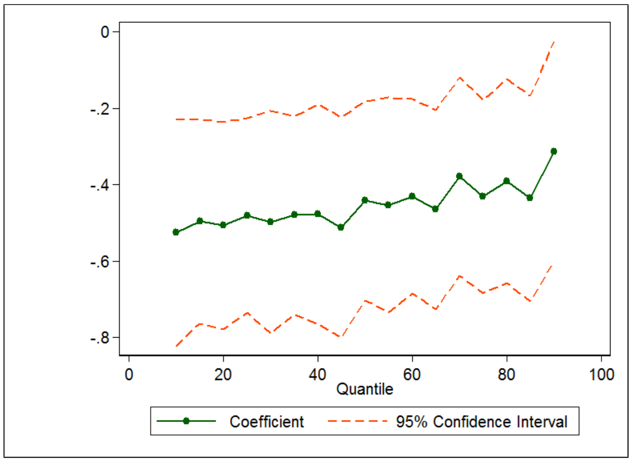
This graph incorporates working-age women only. The figure plots grouped instrumental variables quantile regression estimates of the effect of a 1,000 peso increase in Chinese imports per worker. Point estimates and confidence intervals for each quantile are shown. Units on the vertical axis are log points.

Figure 3.11: The Impact of Chinese Exports on the Conditional Wage Distribution on Manufacturing Workers



This graph incorporates workers in manufacturing only. The figure plots grouped instrumental variables quantile regression estimates of the effect of a 1,000 peso increase in Chinese imports per worker. Point estimates and confidence intervals for each quantile are shown. Units on the vertical axis are log points.

Figure 3.12: The Impact of Chinese Exports on the Conditional Wage Distribution on Non-Manufacturing Workers



This graph incorporates workers outside of manufacturing only. The figure plots grouped instrumental variables quantile regression estimates of the effect of a 1,000 peso increase in Chinese imports per worker. Point estimates and confidence intervals for each quantile are shown. Units on the vertical axis are log points.

3.7 Tables

Table 3.1: Trade Values of Imports from China and the U.S. into Mexico

	China	U.S.
1995	520	53,973
1996	760	67,615
1997	1,289	83,214
1998	1,615	93,307
1999	1,920	105,376
2000	2,878	127,690
2001	4,027	114,060
2002	6,274	106,900
2003	9,400	105,723
2004	14,373	111,262
2005	17,696	118,973
2006	24,438	130,810
2007	29,744	139,931
2008	34,690	151,746
2009	32,529	112,789
2010	45,608	145,450

Trade data come from the UN Comtrade Database. All trade values are in millions of US dollars and are not adjusted for inflation.

Table 3.2: Descriptive statistics

	Mean	St. Dev.
Initial share of male population	47.8	1.8
Initial share of population low-skilled	24.6	14.4
Initial share of population medium-low skilled	25.1	4.9
Initial share of population medium-high skilled	26.9	7.0
Initial share of population high-skilled	23.4	11.5
Initial share of population in agriculture	15.8	20.2
Initial share of female population employed	36.2	9.4
Initial share of employment in manufacturing	19.1	10.7
Change in manufacturing share	-2.4	2.5
Change in log (wage)	0.2	0.2
Change in mean years of schooling	1.3	0.5
Change in import exposure	11.6	8.0
Change in employed share	0.3	4.3

Summary statistics are shown across municipalities in Mexico, using the working-age population as weights. All shares are in percentage points. Initial figures are for 2000, and changes are for the difference between 2010 and 2000. The import exposure measure is in thousands of pesos in 2010 currency. Shares are in percentage points.

Table 3.3: Impact of Chinese import exposure on manufacturing share of working-age population

	OLS (1)	OLS (2)	OLS (3)	First Stage (4)	IV (5)
Import exposure	-0.153***	-0.0362*	-0.00254		-0.0191
	(0.0194)	(0.0195)	(0.0174)		(0.0259)
Manuf share$_{t=2000}$		-0.148***	-0.164***	-0.00475	-0.172***
		(0.00978)	(0.0120)	(0.0282)	(0.0125)
Male share$_{t=2000}$			-0.114***	0.168	-0.117***
			(0.0388)	(0.1154)	(0.0383)
Share low-med educ$_{t=2000}$			0.0209	-0.0156	0.0201
			(0.0173)	(0.0277)	(0.0172)
Share med-high educ$_{t=2000}$			0.0153	0.0510	0.0162
			(0.0181)	(0.0360)	(0.0180)
Share high educ$_{t=2000}$			-0.0217	0.0755*	-0.0270*
			(0.0139)	(0.0439)	(0.0141)
Agricultural share$_{t=2000}$			-0.00870	0.0176	-0.0105
			(0.00686)	(0.01472)	(0.00696)
Fem employment rate$_{t=2000}$			-0.0475***	0.0246	-0.0505***
			(0.00898)	(0.0234)	(0.00943)
Import exposure (IV)				0.567***	
				(0.0354)	
Observations	2,330	2,330	2,330	2,330	2,330
Region Dummies	No	No	Yes	Yes	Yes

The dependent variable is the share of the working-age population in manufacturing (in percentage points) in the first set of columns. All models estimate a regression of the 2000-2010 difference in the dependent variable on the 2000-2010 difference in Chinese import exposure. The first stage regression estimates the impact of Chinese import penetration into Mexico on Chinese import penetration into five other countries (Bolivia, Chile, Colombia, Spain, and Venezuela). Some specifications contain start-of-period covariates (as listed). Models are weighted by the working-age population in 2000. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 3.4: Impact of Chinese import exposure on human capital and employment-related outcomes

	High-skill (1)	Schooling (2)	High-skill manuf. (3)	Employed (4)	Unemployed (5)
Import exposure	-0.0974**	-0.0118**	-0.193***	-0.0445	0.00806
	(0.0410)	(0.00466)	(0.0568)	(0.0305)	(0.00872)
Observations	2,330	2,330	2,313	2,330	2,330
Region Dummies	Yes	Yes	Yes	Yes	Yes

The dependent variable is listed in the column title. All outcomes are calculated as percentage points. All models estimate an instrumental variables regression of the 2000-2010 difference in the dependent variable on the 2000-2010 difference in Chinese import exposure. All specifications contain the full set of start-of-period covariates in column 3 in the first table of results. Models are weighted by the working-age population in 2000. *** p<0.01, ** p<0.05, * p<0.1.

Table 3.5: Impact of Chinese import exposure on wages

	OLS (1)	OLS (2)	OLS (3)	IV (4)
Import exposure	-1.056***	-0.781***	-0.319***	-0.536***
	(0.148)	(0.168)	(0.101)	(0.131)
Manuf share$_{t=2000}$		-0.346***	-0.184**	-0.109
		(0.0757)	(0.0800)	(0.0804)
Male share$_{t=2000}$			0.2798	0.3103
			(0.295)	(0.299)
Share low-med educ$_{t=2000}$			-0.633***	-0.625***
			(0.203)	(0.203)
Share med-high educ$_{t=2000}$			0.257*	0.248*
			(0.132)	(0.134)
Share high educ$_{t=2000}$			-0.384***	-0.331***
			(0.110)	(0.114)
Agricultural share$_{t=2000}$			0.218***	0.236***
			(0.0540)	(0.0538)
Fem employment rate$_{t=2000}$			0.161	0.191*
			(0.103)	(0.105)
Observations	2,330	2,330	2,330	2,330
Region Dummies	No	No	Yes	Yes

The dependent variable is the difference in the average log wage. All models estimate either an OLS or an instrumental variables regression of the 2000-2010 difference in the log wage on the 2000-2010 difference in Chinese import exposure. Some specifications contain start-of-period covariates (as listed). Models are weighted by the working-age population in 2000. *** p<0.01, ** p<0.05, * p<0.1.

Table 3.6: Impact of Chinese import exposure on wages

	All (1)	Men (2)	Women (3)	Manuf. (4)	Non-manuf. (5)
Import exposure	-0.536***	-0.530***	-0.564***	-0.763***	-0.478***
	(0.131)	(0.125)	(0.158)	(0.150)	(0.133)
Observations	2,330	2,330	2,324	2,257	2,330
Region Dummies	Yes	Yes	Yes	Yes	Yes

The dependent variable is the difference in the average log wage for the group listed in the column title. All models estimate an instrumental variables regression of the 2000-2010 difference in the log wage on the 2000-2010 difference in Chinese import exposure. All specifications contain the full set of start-of-period covariates in column 3 in the first table of results. Models are weighted by the working-age population in 2000. *** p<0.01, ** p<0.05, * p<0.1.

Table 3.7: Impact of Chinese import exposure on wages and earnings

	2000-2010	2000-2010	2000-2010	2000-2010	2000-2010	2000-2010	2000-2015
	Original sample	Omitting Mexico City only	Omitting Mexico City metro area	Excluding outliers	Earnings (full-time)	Earnings (all)	Earnings (all)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Import exposure	-0.536***	-0.559***	-0.633***	-0.768**	-0.557***	-0.590***	-0.372***
	(0.131)	(0.136)	(0.147)	(0.336)	(0.126)	(0.129)	(0.0865)
N	2,330	2,324	2,179	2,211	2,330	2,330	2,320
Region Dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes

The dependent variable is the difference in the mean log wage (first four columns) or mean log earnings (last three columns). All models estimate 2SLS regressions of the difference in the dependent variable on the difference in Chinese import exposure between 2000-2010. The last column estimates a model using the difference between 2000 and 2015. All specifications contain the full set of start-of-period covariates in column 3 in the first table of results. Models are weighted by the working-age population in 2000. *** p<0.01, ** p<0.05, * p<0.1.

Table 3.8: Impact of Chinese import exposure on wages (including international exposure)

	OLS (1)	OLS (2)	OLS (3)	IV (4)
Import exposure	-0.600***	-0.461***	-0.227***	-0.257***
	(0.0767)	(0.0883)	(0.0531)	(0.0584)
Manuf share$_{t=2000}$		-0.318***	-0.147*	-0.128*
		(0.0717)	(0.0771)	(0.0730)
Male share$_{t=2000}$			0.252	0.254
			(0.293)	(0.292)
Share low-med educ$_{t=2000}$			-0.620***	-0.617***
			(0.203)	(0.202)
Share med-high educ$_{t=2000}$			0.241*	0.237*
			(0.133)	(0.134)
Share high educ$_{t=2000}$			-0.363***	-0.349***
			(0.109)	(0.112)
Agricultural share$_{t=2000}$			0.228***	0.233***
			(0.0535)	(0.0528)
Fem employment rate$_{t=2000}$			0.179*	0.188*
			(0.103)	(0.102)
Observations	2,330	2,330	2,330	2,330
Region Dummies	No	No	Yes	Yes

The dependent variable is the difference in the average log wage. All models estimate either an OLS or an instrumental variables regression of the 2000-2010 difference in the log wage on the 2000-2010 difference in Chinese import exposure. Some specifications contain start-of-period covariates (as listed). Models are weighted by the working-age population in 2000. *** p<0.01, ** p<0.05, * p<0.1.

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